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"Essays on Exchange Rate and Interest Rate Fluctuations"

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Contents

List of Tables

List of Figures

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Declaration

I declare that any material contained in this thesis has not been submitted for a degree to any other university. I further declare that one paper titled: "New Evidence on the Forward Unbiasedness Hypothesis in the Foreign Exchange Market", drawn from Chapter Two of this thesis has been published in the *Journal of Futures Markets*, 2006, 26(7), 627-656 (co-authored with Prof. Lucio Sarno). Also, the paper "The behaviour of the real exchange rate: Evidence from regression quantiles" drawn from Chapter Three of this thesis appears as Working paper (WP) in the *ECB Working Paper Series (WP 667)* and is under review with the *Journal of Banking and Finance* (revise and resubmit).

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Abstract

The aim of this thesis is to further investigate new empirical methods, results and implications on major topics relating to foreign exchange and interest rate markets. To this end, this thesis is organised in three chapters. The first chapter focuses on nominal exchange rates. It extends the literature of foreign exchange unbiasedness by including information from different derivatives markets. For the purpose of this thesis, it also implicitly provides a lead on the behaviour of interest rate differentials. The second chapter uses innovative econometric methodologies to add new insights in the behaviour of real exchange rates. Finally, Chapter Three explicitly models the international linkages between the interest rate differentials across countries with clear monetary policy implications.

More specifically, a large empirical literature has tested the unbiasedness hypothesis in the foreign exchange market using forward exchange rates. In the first chapter we amend the conventional testing framework to exploit the information in currency options, using a newly constructed data set for three major dollar exchange rates. The main results are that: (i) tests based on stationary regressions suggest that options provide biased predictions of the future spot exchange rate; (ii) cointegration-based tests that are robust to several statistical problems afflicting stationary regressions and allow for endogeneity issues arising from a potential omitted risk premium term are supportive of unbiasedness.

In the second chapter we test for mean reversion in real exchange rates using a recently developed unit root test for non-normal processes based on quantile autoregression inference in semi-parametric and non-parametric settings. The quantile regression approach allows us to directly capture the impact of different magnitudes of shocks that hit the real exchange rate, conditional on its past history, and can detect asymmetric, dynamic adjustment of the real exchange rate towards its long run equilibrium. Our results suggest that large shocks tend to induce strong mean reverting tendencies in the exchange rate, with half lives less than one year in the extreme quantiles. Mean reversion is faster when large shocks originate at points of large real exchange rate deviations from the long run equilibrium. However, in the absence of shocks no mean reversion is observed. Finally, we report asymmetries in the dynamic adjustment of the RER.

Finally, in the third chapter we employ dynamic factor modelling and maximum likelihood estimation to investigate the existence, the patterns and the implications of common fluctuations in the money market rate differentials of a group of countries vis-á-vis the US or Germany. To the extent that money market rates reflect monetary policy decisions we argue that the resulting global factor represents the common part of monetary policy deviations across countries. We find that a significant part of such policy deviations is shared across countries and in fact is mainly driven by the policy interactions of the EU and the US. In particular, the US interest rate seems to emerge as a potential global interest rate. The implication is that policy makers should pay closer attention to foreign policies when setting domestic ones.

Chapter One

1 Introduction

This thesis aims to investigate new empirical methods, results and implications on major topics relating to foreign exchange and interest rate markets. The behaviour of foreign exchange rates and interest rates has attracted a significant amount of attention from both academics and practitioners. This is because fundamental relationships of the kind studied in this thesis form the cornerstone of classical, theoretical models in international finance and, therefore, the base for both academic research and policy decision-making. This thesis provides original extensions to the related literature, adopting always an international perspective and focusing on major industrialised economies. To this end, we investigate the behaviour of nominal exchange rates and re-examine foreign exchange unbiasedness by including information from different derivatives markets. We also use innovative econometric methodologies to add new insights in the behaviour of real exchange rates. Finally, we explicitly model the international linkages between the interest rate differentials across countries with clear monetary policy implications.

More specifically, Chapter Two of this thesis revisits the unbiasedness hypothesis in the context of the foreign exchange (FX) market, one of the most researched and yet controversial hypotheses in the international finance literature. The unbiasedness hypothesis is related to the notion of FX market efficiency, as summarised by the uncovered interest rate parity (UIP) condition, which states that the expected exchange rate change should equal the current interest rate differential–or, in the absence of arbitrage, the forward premium (the difference between the forward and spot rates). Under UIP and in the absence of arbitrage (i.e. assuming that covered interest parity holds), the forward exchange rate provides an unbiased forecast of the future spot exchange rate, or, equivalently, the forward premium provides an unbiased forecast of the future change in the spot exchange rate–this is the key assertion of the unbiasedness hypothesis.

However, attempts to empirically justify the theoretical relationship between changes in the exchange rate and the forward premium were met with less than satisfactory results. In a highly cited paper, Fama (1984) suggests that the expected change in the exchange rate is often inversely related to the forward premium, in stark contrast with UIP. This realisation has spurred an enormous amount of research and produced a large spectrum of results, which gave way to an extended list of possible explanations (e.g. Lewis, 1995; Flood and Rose, 1996; Engel, 1996; Chinn and Meredith, 2004; Sarno, 2005). In general, however, tests of the unbiasedness hypothesis for different currency pairs and time periods gave further credit to Fama's results, which are now considered a stylised fact (Froot and Thaler, 1990), giving rise to the "forward bias puzzle," one of the central puzzles in international finance.

In this thesis the unbiasedness hypothesis is re-examined, by changing vehicles of forming predictions about the future spot exchange rate. Specifically, the focus switches from the forward to the options market. Using data from the Philadelphia Exchange (PHLX), we construct a synthetic forward contract, made of currency options, which we call "option equivalent contract" and substitute it for the standard forward contract in the analysis of unbiasedness. These results are compared with the results obtained using forward contracts. Throughout this study, conventional methods (the typical UIP regression first used by Fama) are combined with cointegration-based tests for unbiasedness, so as to present a thorough analysis for both forward- and option-based unbiasedness tests.

This approach yields several original additions to the relevant literature. First, evidence is provided on the empirical validity of the unbiasedness hypothesis using currency options in a novel testing framework, complementing the conventional testing procedure which was, until now, restricted to forward markets. Second, our unique data set allows us to create a bridge between the over-the-counter (OTC) market for forward contracts and the organised exchange for options contracts. This research directly compares the two derivatives markets, the forward and the options market, in terms of the statistical properties of the resulting contracts. Third, the empirical work allows us to assess whether the bias puzzle recorded in the literature to date is "forward specific" or a problem of a more general nature which is likely to be pervasive in other derivatives contracts.

The results in this thesis provide several useful insights. The methodology used to produce the option equivalent contracts results into a synthetic forward which may be compared to the conventional forward contract, from which it differs in terms of contract specifications (e.g. maturity, expiry, trading specifications). The resulting option equivalent and the forward rate exhibit striking similarities in terms of both statistical properties and test results relating to the unbiasedness hypothesis. Overall, conventional tests based on stationary regressions suggest that there is an "option bias puzzle," reinforcing the case for the well documented "forward bias puzzle." However, more powerful cointegration tests that are robust to several statistical problems afflicting stationary regressions and allow for the presence of a potential risk premium–which does not, *per se*, preclude unbiasedness–show ample support for unbiasedness for both forwards and options. The results are found to be robust to a variety of different departures from the core analysis, including the frequency of the data and the maturity of the derivatives contract.

The third chapter of this thesis concerns the notion of the Purchasing Power Parity (PPP). The PPP suggests that, in the absence of arbitrage, aggregate price levels of tradeable goods between two countries should be equal, if expressed in the same currency. It is the simplest model of exchange rate determination and a fundamental arbitrage relationship in international asset pricing. In that sense, PPP provides an equilibrium relationship for the real exchange rate (RER), which is the nominal exchange rate, adjusted for relative price levels. If PPP holds, the relative price levels and/or the bilateral nominal exchange rate would adjust in such a way so that the RER remain constant. In that sense, variations in the RER would suggest deviations from PPP.

Although intuitive theoretically, PPP lacks strong empirical support. In practice, the RER exhibits high variability over time and spends long periods away from its suggested PPP equilibrium. The ambiguity surrounding the persistency of the RER and the validity of PPP were evident in the early empirical attempts, which were clearly rejecting PPP. Nevertheless, it soon became obvious that a potential reason for this apparent failure was the underlying assumption of constant dynamics for the RER process. More specifically, the speed of adjustment towards the PPP equilibrium would be constant, no matter how far the RER is from its equilibrium value, or how big is the shock that hits the RER.

This belief was soon challenged by both theoretical and empirical research, which relied on arbitrage arguments to show that the speed of adjustment should vary according to the degree of disequilibrium from PPP (see amongst others Dumas, 1992; Obstfeld and Taylor, 1997; Taylor, Peel and Sarno, 2001). Allowing for the RER to assume different speeds of adjustment spurred a long literature of increasingly sophisticated RER models, frequently relying on competing theoretical arguments about the causes and the specification of the dynamic adjustment of the RER. Empirical evidence from this literature provides evidence in favour of PPP and shows that the speed of adjustment is much faster when the RER is far away from its PPP equilibrium, or when the RER is hit by large shocks. Under this reasoning, the early tests were only capturing an average behaviour of the RER and were therefore suffering from an upward bias when measuring the persistence of the RER.

Another relevant issue in the PPP literature concerns the potentially different speeds of adjustment for positive or negative deviations of the RER from its PPP equilibrium, i.e. the possibility of asymmetric mean reversion towards the RER equilibrium. Theoretical and empirical arguments can be found in support to both sides. On the one hand, if goods arbitrage drives the impetus back towards the PPP equilibrium, the speed of adjustment needs not be different above or below the equilibrium (Taylor, Peel and Sarno, 2001). On the other hand, a more recent strand of literature brings forward the role of central bank intervention as an underlying force affecting the dynamic adjustment of the RER. In this context, asymmetries may arise as a result of intervention policies directed at the RER (Dutta and Leon, 2002; Leon and Najarian, 2005).

In this chapter, we investigate the RER behaviour using the recently developed method of quantile regressions in semi-parametric and non-parametric settings. This methodology does not focus on the conditional mean function (i.e. does not assume constant speed of adjustment), but rather studies a whole range of conditional quantile functions. In so doing, the quantile approach potentially encapsulates assumptions and results from previous less-general, parametric models, in an a-theoretical and flexible way. It thus circumvents the need to discriminate across different formulations of the causes and specification of the dynamic adjustment of the RER to its long run equilibrium.

We use this method as an alternative framework for testing the dynamic adjustment of the RER. To this end we investigate potentially different adjustment speeds when different magnitudes of shocks hit the RER. Different solutions at distinct quantiles may be interpreted as differences in the mean reverting behaviour of the RER at different magnitudes of RER shocks. We further investigate how this effect changes with initial conditions (i.e. the degree of disequilibrium of the RER from its PPP value when the shock hits the RER). Moreover, the quantile method is able to reveal asymmetries in both the distribution of RER shocks and their impact on the RER mean reverting behaviour in a simple, intuitive and yet effective way. In short, the quantile framework is modified to incorporate the (potentially asymmetric) effects of various sizes of RER shocks, and is, therefore, a more robust PPP test compared to standard models.

More specifically, our results suggest that: a) the dynamic behaviour of the RER is affected by the magnitude of RER shocks, with large RER shocks inducing potentially strong mean reverting tendencies. b) When large shocks to the RER originate at large RER disequilibrium levels (i.e. far away from its PPP equilibrium), the effect can be even stronger. c) On the contrary, small shocks to the RER considerably weaken mean reversion tendencies, irrespective of the disequilibrium point of the RER at the time of the shock. d) There are marked asymmetries in the behaviour of the RER, i.e. extreme positive shocks (depreciations) can generate different reversion patterns than extreme negative shocks (appreciations). Their extent also depends on the original condition of the RER with respect to its long run equilibrium. Overall, the quantile analysis provides original insights into the PPP literature but also further refines and enhances results obtained by the up-to-date relevant research.

Having investigated some major aspects of the nominal and real exchange rate, i.e. the opportunity cost of holding money *of* one country rather than another, we now turn our attention to the opportunity cost of holding money *in* one country rather than another, that is, the interest rate differential. Always adopting an international perspective, we study the international linkages between interest rate differentials as a final topic (Chapter Four) in this thesis. Our motivation draws mainly from previous influential work exposing highly synchronised fluctuations in fundamental economic variables across countries. Namely, a wide literature has provided evidence on the existence of strong comovements in business cycles, fostering the idea of a world business cycle (see Kose *et. al.*, 2004 and the references therein). At the same time evidence has also emerged of a world inflation rate (Mojon and Ciccareli, 2005). It is therefore expected that such strong international linkages in output and inflation should also induce strong comovements in monetary policies across countries and therefore, short term interest rates. Moreover, if we accept that certain countries are considered as global "policy setters", it is possible to observe a certain degree of homogeneity or comovement in the response of the monetary policies of a group of countries vis-á-vis these global players, as this is expressed in money market differentials. Furthermore, it could be that there is a specific pattern in the responses of the rest of the world to the "policy setter", for example the common reaction can be dominated by the reaction of a specific country vis-á-vis the "policy setter" country.

Close to this line of thinking, previous research has reported high sensitivity of local interest rates to international ones (see for example Frankel *et.al.*, 2004; Diebold *et.al.*, 2006 and Dungey *et.al.*, 2000) and also suggested the existence of a world interest rate, most probably driven by the US rates (Chinn and Frankel, 2005). Nevertheless, a concise examination for the existence, the nature and the implications of a common factor in short term interest rate differentials across countries was missing from the literature and this is the gap that our contribution is aiming to fill.

We use dynamic factor modelling and maximum likelihood estimation techniques in an effort to investigate the common fluctuations in the money market rate differentials of a group of major countries vis-á-vis a common denominating country. We follow the intuition of Frankel *et.al.* (2004) and Chinn and Frankel (2005) and use the US and Germany as the denominating countries. In our setting the interest rate differentials of each country are explained by a common/global factor and an orthogonal idiosyncratic, domestic component. Our assumption is that the monetary policy stance of each country is reflected in the money-market interest rates, that is, interest rates with maturities up to one year. In that sense money market rates at different maturities reflect the different degrees of infiltration of monetary policy into the domestic money market rates at different horizons. Following this line of reasoning, the global factor captures the common fluctuations in the monetary policy deviations of the countries under consideration with respect to the denominating country. Therefore, our study focuses not on the similarities but on the discrepancies between the monetary policies of our group of countries vis-á-vis the US (German) policies. By exploring the extent to which these discrepancies are driven by a common force, we provide further evidence of integration in a globalised environment.

We aim to reveal valuable insights in the behaviour of the short-term interest rate differentials with direct implications for monetary policy actions. More specifically, we would like to measure the extent of interest rate differentials comovements across countries and identify countries with global status, in the sense that they affect the behaviour of the global factor. In this direction, our methodology allows us the advantage to investigate finer interactions among global policy makers, namely, whether the policy divergencies vis-á-vis the denominating country are driven by the behaviour of a specific country. Moreover, we further investigate whether global players can act as interest rate setters, that is, set their interest rates independently. Finally, we draw inferences on the existence of a world interest rate.

We report evidence on strong comovements in the interest rate differentials of a group of countries vis-á-vis the denominating country. In other words, there seems to be a strong global factor driving the interest rate differentials of several countries, suggesting increased sensitivity of the domestic interest rate differentials to the global one. However it is notable that sensitivity patterns might change depending on the denominating country. Nevertheless, the global factor seems to be reacting to monetary policies of the US and the EU. In fact, it appears that when the US is the denominating country, the global factor seems to be driven primarily by EU policies vis-á-vis the US. Nevertheless, the Euro-area and the US seem to enjoy a certain degree of independence when choosing their own policies. Finally, there seems to be evidence that the US policy rate emerges as a global interest rate.

In terms of monetary policy implications, our results suggest that although there are still discrepancies between the monetary policies of various countries, these discrepancies are not erratic, but rather coordinated and influenced by major global players like the EU and the US. Therefore, monetary policy-makers should pay closer attention to foreign macroeconomic aggregates, and the resulting monetary policy choices of global players.

The remainder of this thesis is set up as follows: Chapter Two describes the notion of unbiasedness in the foreign exchange market, introduces the option equivalent to the forward contract and presents results from the empirical comparison of the two types of contracts. Chapter Three exposes the previous PPP literature and presents the advantages of the quantile regression analysis in the subsequent empirical analysis. Chapter Four explains the intuition behind the existence, the meaning and the implications of a global factor driving international interest rate differentials. The subsequent empirical analysis draws clear policy implications. Finally, Chapter Five briefly summarises the key findings of this thesis and concludes.

Chapter Two

2 Revisiting the Forward Unbiasedness Hypothesis in the Foreign Exchange Market: Evidence from options markets

2.1 Introduction

This chapter revisits the unbiasedness hypothesis in the context of the foreign exchange (FX) market, one of the most researched and yet controversial hypotheses in the international finance literature. The unbiasedness hypothesis is related to the notion of FX market efficiency, as summarised by the uncovered interest rate parity (UIP) condition, which states that the expected exchange rate change should equal the current interest rate differential or, in the absence of arbitrage, the forward premium (the difference between the forward and spot rates). Under UIP and in the absence of arbitrage (i.e. assuming that covered interest parity holds), the forward exchange rate provides an unbiased forecast of the future spot exchange rate, or, equivalently, the forward premium provides an unbiased forecast of the future change in the spot exchange rate–this is the key assertion of the unbiasedness hypothesis[1].

The profession has long focused on investigating the relationship between changes in the exchange rate and the forward premium with less than satisfactory results. In a highly cited chapter, Fama (1984) suggests that the expected change in the exchange rate is often inversely related to the forward premium, in stark contrast with UIP. This realisation has spurred an enormous amount of research and produced a large spectrum

[1]In our terminology in this section, tests of UIP are essentially interchangeable with tests of the unbiasedness hypothesis–that is the coefficient on lagged interest differentials or forward premia (lagged forward rate) in regressions of current exchange rate changes (current exchange rate level) is unity. This is somewhat loose, in that UIP is a sufficient but not necessary condition for unbiasedness, as discussed later in the paper.

of results, which gave way to an extended list of possible explanations (e.g. Lewis, 1995; Flood and Rose, 1996; Engel, 1996; Chinn and Meredith, 2004; Sarno, 2005). In general, however, tests of the unbiasedness hypothesis for different currency pairs and time periods gave further credit to Fama's results, which are now considered a stylised fact (Froot and Thaler, 1990), giving rise to the "forward bias puzzle," one of the central puzzles in international finance.

In this chapter the unbiasedness hypothesis is re-examined, by changing vehicles of forming predictions about the future spot exchange rate. Specifically, the focus switches from the forward to the options market. Using data from the Philadelphia Exchange (PHLX), we construct a synthetic forward contract, made of currency options, which we call "option equivalent contract" and substitute it for the standard forward contract in the analysis of unbiasedness. These results are compared with the results obtained using forward contracts. Throughout this study, conventional methods (the typical UIP regression first used by Fama) are combined with cointegration-based tests for unbiasedness, so as to present a thorough analysis for both forward- and option-based unbiasedness tests.

This approach yields several original additions to the relevant literature. First, evidence is provided on the empirical validity of the unbiasedness hypothesis using currency options in a novel testing framework, complementing the conventional testing procedure which was, until now, restricted to forward markets. Second, our unique data set allows us to create a bridge between the over-the-counter (OTC) market for forward contracts and the organised exchange for options contracts. This research directly compares the two derivatives markets, the forward and the options market, in terms of the statistical properties of the resulting contracts. Third, the empirical work allows us to assess whether the bias puzzle recorded in the literature to date is "forward specific" or a problem of a more general nature which is likely to be pervasive in other derivatives contracts[2].

[2]In other words, we address the question whether the puzzle is likely to be caused by some specific

The results in this chapter provide several useful insights. The methodology used to produce the option equivalent contracts results into a synthetic forward which may be compared to the conventional forward contract, from which it differs in terms of contract specifications (e.g. maturity, expiry, trading specifications). The resulting option equivalent and the forward rate exhibit striking similarities in terms of both statistical properties and test results relating to the unbiasedness hypothesis. Overall, conventional tests based on stationary regressions suggest that there is an "option bias puzzle," reinforcing the case for the well documented "forward bias puzzle." However, more powerful cointegration tests that are robust to several statistical problems afflicting stationary regressions and allow for the presence of a potential risk premium–which does not, *per se*, preclude unbiasedness–show ample support for unbiasedness for both forwards and options. The results are found to be robust to a variety of different departures from the core analysis, including the frequency of the data and the maturity of the derivatives contract.

The structure of the chapter is as follows. Section 2.2 presents a brief review of the literature on testing the unbiasedness hypothesis and the motivation for the use of options in this context. Section 2.3 describes the data set and provides details related to the construction of the synthetic forward. Section 2.4 presents the core empirical results, while Section 2.5 reports robustness checks of the core results. Section 2.6 briefly summarises and concludes.

2.2 Testing the unbiasedness hypothesis using options

This section briefly reviews the enormous literature testing the validity of UIP and the unbiasedness hypothesis in the FX market, which has led to mixed results. Specifically, on the one hand tests based on stationary regressions (e.g. research following Fama, 1984) have recorded that the forward premium is not an unbiased predictor of the

characteristics of the forward market (e.g. the specific way that agents in this market form predictions about the future spot rate), or it is pervasive in other derivatives markets as well.

future rate of depreciation, and in fact there is a forward bias such that the forward premium is generally inversely related to future movements in the exchange rate. On the other hand, more recent cointegration-based tests provide some supportive evidence for the forward rate unbiasedness hypothesis (Barnhart, McNown and Wallace, 1999; Maynard, 2003), although cointegration studies provide, overall, mixed results[3]. We then describe how the conventional unbiasedness tests are amended by substituting the forward exchange rate with a suitably constructed proxy for the market expectation, based on information embedded in options contracts. Such proxy is termed the "option equivalent."

2.2.1 Conventional tests of forward rate unbiasedness

UIP purports that the FX gain from holding one currency instead of another–the expected exchange rate change–must be offset by the opportunity cost of holding funds in one currency rather than the other–the interest rate differential:

$$\Delta_k s^e_{t+k} = i_{t,k} - i^*_{t,k} \tag{2.1}$$

where s_t denotes the logarithm of the spot exchange rate (domestic price of foreign currency) at time t; i_t and i^*_t are the nominal interest rates available on similar domestic and foreign securities respectively (with k periods to maturity); $\Delta_k s^e_{t+k} \equiv s^e_{t+k} - s_t$; and the superscript e denotes the market expectation based on information at time t. In its simplest form, FX market efficiency can be reduced to a joint hypothesis that FX market participants are, in an aggregate sense, (a) endowed with rational expectations and (b) risk-neutral. The hypothesis can be modified to adjust for risk, so that it then becomes a joint hypothesis of a model of equilibrium returns (which may admit risk

[3]The lack of consensus in empirical research on forward rate unbiasedness is well characterised by Engel (1996, p. 141) as follows: 'To summarise [...] some have found [the future exchange rate and the current forward rate] are cointegrated with cointegrating vector $(1, -1)$; some have found they are cointegrated but not with a cointegrating vector $(1, -1)$; and some have found that they are not cointegrated. These conflicting results hold on tests for the same set of currencies.

premia) and rational expectations.

In practice researchers investigate UIP with the aid of Covered Interest Parity (CIP), the most common no-arbitrage relationship in the context of the FX market. CIP finds its mathematical representation in the form: $f_t^k - s_t = i_{t,k} - i_{t,k}^*$, where f_t^k is the logarithm of the k-period forward rate (i.e. the rate agreed now for an exchange of currencies k periods ahead). Should CIP not hold at a point in time, profitable opportunities would emerge, which would induce trade in opposite directions resulting to their elimination[4].

Assuming that CIP holds, UIP can be re-written as $\Delta_k s_{t+k}^e = f_t^k - s_t$, i.e. the forward premium (or forward discount) $f_t^k - s_t$ should equal the market expectation of the exchange rate change $\Delta_k s_{t+k}^e$; and $s_{t+k}^e = f_t^k$, i.e. the forward rate should be an unbiased predictor of the future spot rate. A test of this hypothesis involves regressing the exchange rate change on the lagged forward premium and, following much previous literature, we shall refer to this regression as the 'Fama regression' (Fama, 1984):

$$\Delta_k s_{t+k} = \alpha + \beta \left(f_t^k - s_t \right) + \varepsilon_{t+k} \tag{2.2}$$

where, under UIP, $\alpha = 0$, $\beta = 1$, and ε_{t+k} is a white noise error. The empirical results from estimating regression (2.2) have led to strong rejections of UIP and, hence, FX market efficiency (e.g. see the references in the survey of Hodrick, 1987; Lewis, 1995; Engel, 1996). While α is generally close to zero and often statistically insignificant, β is estimated to be far from its theoretical value of unity and it is often found to be negative and statistically significantly different from zero. Indeed, it is a stylised fact that estimates of the slope parameter β are generally closer to minus unity rather than plus unity (Froot and Thaler, 1990). The negative value of β is the central feature of the forward bias puzzle, one of the most robust puzzles in international finance, which

[4]Extensive empirical evidence provides support to the validity of CIP (for a survey of this evidence, see e.g. Sarno and Taylor, 2003, Ch. 2). Note that, unlike CIP, UIP is not an arbitrage condition since one of the terms in the UIP equation, namely the exchange rate at time $t + k$, is unknown at time t and, therefore, non-zero deviations from UIP do not necessarily imply the existence of arbitrage profits due to the FX risk associated with future exchange rate movements.

remains unexplained even with 20 years of hindsight since the work of Fama (1984)[5].

One potential problem with estimation of β in the Fama regression in equation (2.2) is the unbalanced nature of the regression that arises if the forward premium is a nonstationary or fractionally integrated process, while the exchange rate change is stationary, highly volatile and near white noise (e.g. Baillie and Bollerslev, 1994). In this case, the unbalanced nature of the Fama regression confounds statistical inference and makes it very difficult to accurately estimate β. The problem becomes even more severe when considering that, in addition to the strong persistence of the forward premium, the exchange rate displays very persistent volatility. Notably, Baillie and Bollerslev (2004) show that a model calibrated on realistic parameter values for daily exchange rates which display such volatility patterns and a persistent forward premium would make the convergence of ordinary least squares (OLS) estimates of β to the true β very slow even when UIP holds and the true value of β is unity. In turn, this makes it difficult to test unbiasedness on the basis of the conventional Fama regression.

Another strand of the literature, building on ideas initially put forth by Fama (1984), and further elaborated by Liu and Maddala (1992) and Barnhart, McNown and Wallace (1999), claims that the conventional Fama regression is invalidated, due to problems of endogeneity, which may result from the appearance of an unobserved risk premium. Specifically, note that the vast majority of studies in this context estimate the Fama regression using OLS. This can be problematic in the presence of an omitted risk premium in the Fama regression, in which case OLS would yield biased and inconsistent estimates of β due to a simultaneity problem (Fama, 1984; Liu and Maddala, 1992; McCallum, 1994). Recently, Barnhart, McNown and Wallace (1999) have formally shown that two conditions are needed for the simultaneity problem to arise: (i) the forward rate must be a function of an unobservable omitted variable, such as predictable

[5]Exceptions include Bansal and Dahlquist (2000), who document that the forward bias is largely confined to developed economies and to countries for which the US interest rate exceeds foreign interest rates; and Flood and Rose (2002), who report that the failure of UIP is less severe during the 1990s and for countries which have faced currency crises over the sample period investigated.

excess returns; (ii) the term containing the forward rate in the estimated regression must be stationary or, if nonstationary, can be normalised to a stationary variable. Under these conditions, Barnhart, McNown and Wallace document the severity of this problem in a variety of spot-forward regressions, concluding that most common tests of unbiasedness are non-informative in the presence of simultaneity. Failure to properly account for these factors results into correlation between the forward premium and the error term, which induces the bias to assume values bigger than unity, therefore driving the OLS estimate of β towards negative values. This simultaneity problem renders the estimates from the Fama regression, and other derivative formulations of UIP tests, biased and inconsistent.

The proposed remedial methodology is to carry out a cointegration analysis involving the level of the spot exchange rate and the lagged forward rate (an unbiased predictor of spot rate under the unbiasedness hypothesis). Specifically, Barnhart, McNown and Wallace (1999) formally demonstrate that a forward unbiasedness test that is immune from the endogeneity problem involves two steps: first, testing for the existence of a cointegrating relationship of form $[1, -1]$ between s_{t+k} and f_t^k; second, if this cointegrating relationship holds, then a test for forward unbiasedness involves testing for residual correlation (both own and cross-currency correlation). Forward unbiasedness holds if this exact cointegrating relationship is validated by the data *and* the stationary residuals are white noise, suggesting that no incremental information can be added using available information at time t. In contrast to the results from estimating the Fama regression (2.2), the evidence from forward unbiasedness tests based on a cointegrating framework, which effectively allows for a potential risk premium term in the relationship between forward rates and future spot rates, and subsequent residual tests lends some support to the hypothesis that the forward rate is an unbiased predictor of the future spot exchange rate (e.g. Liu and Maddala, 1992; Barnhart, McNown and Wallace, 1999)[6].

[6]For interesting related results in the context of cointegration in other derivatives markets, see Kellard, Newbold, Rayner and Ennew (1999).

It is worthwhile noting that tests based on cointegration are also robust to the problems induced by the persistence in the forward premium and in the volatility of exchange rate changes, since the cointegrating regression in levels does not require strong conditions on the behaviour of the error term. Essentially the only condition required is stationarity of the error term for cointegration to be established and for obtaining a consistent estimate of β[7]. Therefore, there are at least two reasons why cointegration tests may be more favourable to forward unbiasedness: the presence of an omitted risk premium and the potentially unbalanced nature of the Fama regression due to the persistence properties of the forward premium and of exchange rate volatility. Cointegration tests do not allow us to disentangle between these two issues, but they do allow us to provide a more robust test of unbiasedness.

In essence, the literature provides mixed evidence on the validity of the forward unbiasedness hypothesis. Studies employing the Fama regression provide robust evidence of a forward bias puzzle (β different from unity and often negative or statistically insignificant), whereas some more recent studies based on a cointegrating framework suggest that the forward rate is unbiased predictor of the future spot rate. All of this evidence is based, however, on one specific derivatives contract, namely the forward exchange rate, in order to proxy the market expectation of the future spot rate.

2.2.2 Using options to test the unbiasedness hypothesis

In this chapter we endeavour to find a different and yet simple path to test the unbiasedness hypothesis. Our shift of focus on FX options, as predictors of the future spot exchange rate, not only bears some plausible intuition, but also acts as a robustness check for the previous results documented in the literature based on forward contracts.

The reasoning for choosing the options market resides in our effort to extract infor-

[7] The precise definition of cointegration requires the cointegrating vector to be covariance stationary. Hansen (1992) shows that much of the statistical theory developed under the strict definition of cointegration still holds when heteroskedasticity is permitted in the cointegrating vector.

mation from a different FX derivatives market than the conventional forward market, yet bearing an extensive involvement in the FX market practices. Apart from trading and contract setting conventions, which induce differences between the forward and options markets (briefly presented in Section 2.3), two intuitions are offered as to why options may contain somewhat different information from forwards. In analysing options, one should bear in mind that, intuitively, an option contract represents a bet that the price of the currency examined will be above or below a certain level. Investors who believe that the price will rise buy call options and those who believe that the price will fall buy put options. We are interested in investigating whether the different tenets of the two markets, as described below, would induce or compel different betting behaviour by agents–see Breuer and Wohar (1996) for a discussion of the specific institutional features of forward markets and their relevance to tests of unbiasedness.

First, it may seem that options contain different information than forward contracts due to the flexibility (of whether and/or when) to exercise an option[8]. In the case of options, the strike price is a mere reference point as the investor targets a wider range of values above or below the actual strike price. On the contrary, in forward contracts, where no such flexibility exists, the settlement price is the exact betting price and the investor aims for a final result as close to that price as possible.

Second, option contracts at a specific time t can have various strike prices: as the degree of moneyness of the contract changes, contracts with new strike prices are introduced to always ensure the presence of put and call contracts. Further to that, tailor-made contracts are also introduced for different strike prices according to the needs of the clients. Therefore, it is possible to have a whole distribution of strike prices for otherwise identical contracts at each time t, which can potentially better capture the expectations of the market. This again contrasts with forward contracts, for

[8]American options contracts are perhaps the clearest example, because they include both the downward insurance to the investor that European contracts have (in case of an unfavorable outcome the investor leaves the option unexercised and only loses the premium paid to acquire the option) plus an additional time value parameter, which comes from the flexibility to exercise on or before expiry (the investor has the advantage of being able to wait for the most appropriate time to exercise the option).

which researchers are only presented with a single forward rate at time t from available databases. This property adds a further dimension to option contracts, namely the distribution of strike prices, which adds to the widely used term structure of contract prices.

The above traits present us the opportunity to test the unbiasedness hypothesis by applying a different instrument. A synthetic forward contract is created (termed the options equivalent, or simply o). The aim is to re-examine forward unbiasedness by: (a) using all the relevant conventional methods based on both the Fama regression and on cointegration analysis; (b) presenting a thorough investigation of both forward and options by undertaking several robustness tests to investigate whether the two derivatives markets yield similar results in terms of portraying investors' expectations; (c) determining whether the forward bias puzzle is indeed specific to the forward market or a more general feature of FX markets.

2.2.3 Calculating the synthetic forward

In order to compare the forward and options markets, we begin from calculating an option measure that is equivalent to the forward rate, i.e. a synthetic forward contract. To this end, the arbitrage conditions of both the forward and the options market are combined. The most prominent arbitrage condition in the options market is the Put-Call Parity (PCP) condition, which establishes a relationship between European put and call option prices (e.g. Stoll, 1969; Merton, 1973; for the case of options in the FX market see Grabbe, 1983). More specifically, using capital letters to relate to levels (i.e. no longer logs of values), PCP suggests that buying a call $(-C)$ and selling a put $(+P)$ with the same strike price (K) and for the same underlying asset (in our case an exchange rate (S)) must yield exactly the same payoff to an investor as a synthetic long forward; i.e., given a domestic interest rate i and a foreign interest rate i^*, the strategy involves borrowing $K/(1+i)$ domestic, purchasing foreign currency and investing $S/(1+i^*)$ worth of foreign currency abroad (Levich, 2001). Therefore:

$$-C+P=\frac{K}{1+i}-\frac{S}{1+i^*} \tag{2.3}$$

or, equivalently

$$C-P=\frac{S}{1+i^*}-\frac{K}{1+i} \tag{2.4}$$

for the reverse strategy.

Although only a thin branch of the literature has tested the validity of this no-arbitrage condition, the available empirical evidence suggests that observed violations occur rarely and do not last long, supporting the assumption of no arbitrage (Shastri and Tandon, 1985, 1986; Bodurtha and Courtadon, 1986, 1987; El-Mekkaoui and Flood, 1998). CIP and PCP can be combined to get the so called Put-Call Forward (PCF) parity relation (Grabbe, 1983):

$$\frac{F-K}{1+i^*}=(C-P). \tag{2.5}$$

A simple reparametrisation of this parity relationship gives us the desired synthetic forward contract price, termed the option equivalent O, which is essentially a synthetic forward contract made of options:

$$O\equiv F=K+(C-P)(1+i^*). \tag{2.6}$$

This representation relates the forward price to the price of put and call option contracts on the same strike price. Note that this specific formulation applies to European options, while the relevant equation for the option equivalent would hold with inequality for American options.

Intuitively, equation (2.6) suggests that the option equivalent to the forward rate is the amount by which the final price will either exceed (if the call contract is exercised) or fall below (if the put contract is exercised) the strike price. Further details on the construction of the option equivalent are provided in the following section. Defining o as the log-option equivalent, the Fama regression may be written in terms of a link between the spot rate change and the option premium as follows:

$$\Delta_k s_{t+k} = \alpha + \beta \left(o_t^k - s_t\right) + \eta_{t+k} \tag{2.7}$$

which has the usual interpretation, i.e. $\alpha = 0$, $\beta = 1$ and η_{t+k} is a white noise error under FX market efficiency. Note that in equation (2.7) α and β are used to denote the constant term and the slope parameter respectively to ease the comparison with the corresponding parameters in the conventional Fama regression[9] (2.2).

2.3 Data

2.3.1 Sources

The data set employed in the empirical work consists of weekly spot, forward, synthetic forward and interest rate (eurocurrency) data. These data are at weekly frequency, which had to be carefully constructed from intraday data. The sample period spans from 3 January 1986 to 31 December 2003.

The synthetic forward was constructed from intraday data on options, provided by the Philadelphia Exchange (PHLX), the main currency options exchange in the US and the only organised exchange in the US where it is possible to trade currency options on spot exchange rates[10]. The difference between the OTC market and the organised exchanges is significant. OTC markets are decentralised markets that provide flexibility in option contracts (tailor-made contracts), customising their specifications. In contrast,

[9] Clearly, in the case of American options, one would be dealing with an inequality in equation (2.3) for PCP and the resulting option equivalent in equation (2.6) would, therefore, also hold with inequality.

[10] It is worth noting that it is possible to trade currency options on the Chicago Mercantile Exchange (CME). However, trading at the CME is only for options on futures, not on spot exchange rates.

organised exchanges are largely centralised markets, offering standardised contract specifications and market conventions. Organised exchanges represent a smaller venue for currency option trading, compared to the OTC market, although the former is growing at a somewhat stronger pace[11]. Trading in currency options is substantial and has experienced steady growth since the beginning of trading at the PHLX in the early 1980s. As it is well documented (e.g. BIS, 2004), derivatives trading dominates spot trading in terms of volume in currency markets. Forward instruments (outright forwards and swaps) are in a dominant position relative to options transactions, but both segments account for a substantial share of currency derivatives trading. The PHLX is in a dominant position with respect to options trading in the US, being the only organised exchange for currency options on spot rates, although trading activity has declined in the late 1990s and early 2000s. Such decline is not confined to options trading, affecting the FX market more generally[12]. However, the most recent estimates of trading volume in the FX market at the BIS indicate a substantial increase in turnover in the market and in the growth of activity in all segments of currency derivatives instruments.

The PHLX has kindly made available to us the full trading history tape, which consists of all recorded transactions on standardised currency options contracts from 1986 to 2003. Specifically, the tape records the characteristics of all put and call options being traded (underlying currency, option premium, strike price, expiry date and number of contracts trading at the specific price) as well as the spot price of the underlying currency with time precision to the nearest second. Bid and ask spreads for both spot and option prices are being recorded non-continuously and are, therefore, not being used for reasons of consistency within the same series and with the forward data. Thus, the mid-point was used for both spot and option prices, which is consistent

[11] Information on most aspects of interest related to trading in spot and derivatives instruments in the foreign exchange market is available on the 2004 Bank for International Settlements (BIS) *Survey on Foreign Exchange and Derivatives Market Activity.*

[12] Possible causes put forth include the advent of the euro, the consolidation of the banking industry, the growth of electronic trading, and the "events of 1998", characterised by higher risk aversion and a global withdrawal of liquidity (BIS, 2004).

with the literature testing the forward unbiasedness hypothesis. The data used are from 02:30am to 02:30pm (Philadelphia time), which includes the hours of main trading activity throughout the sample[13]. We focus on the most actively traded contracts, namely American contracts with mid-month expiry, for three major dollar exchange rates against the UK sterling, the Japanese yen and the Swiss franc (GBP, JPY and CHF).

Although the PHLX also trades European contracts, American contracts are investigated because they are by far the most heavily traded. We include all available trades, without excluding cases of potential early exercise. Some authors adopt the practice of excluding such cases in the context of testing the validity of PCP (Shastri and Tandon, 1985, 1986; Bodurtha and Courtadon, 1986, 1987). However, we prefer using all of the observations because our focus is not on testing PCP and because excluding trades that are deep in the money would distort the intraday distribution from which our time series are constructed since it would skew it to the left.

It is also worth noting that, although researchers have studied subsets of the PHLX data in previous chapters for other research purposes, the current chapter analyses the longest PHLX span ever considered in empirical work, with the full tape consisting of about 1,800,000 intraday observations for each time series examined.

Data on 1- and 3-month forward contracts, spot rates and interest rates (eurocurrency rates), at the daily frequency, for the same set of currencies and sample periods as above, were provided by the BIS. These data were converted to weekly frequency, in such a way as to match the dates available for our weekly forward equivalent series.

[13]During the history of currency options, which only begun in 1982, the PHLX has experimented with various timing schedules for its operations, in response to demand from different world sectors, resulting to an around-the-clock trading session in 1990. However, lately its operations have been scaled back and its current currency option trading hours are from 02:30am to 02:30pm (Philadelphia time).

2.3.2 Data details and manipulation

Our aim is to transform the intraday data on options contracts into a weekly series of synthetic forward contracts. We focus on a specific day of the week, namely Friday (the day of the contracts' expiry), thus creating weekly time series where each weekly observation corresponds to the last trading day of the week[14]. The intraday option equivalent was constructed using equation (2.6) for the intraday data on each Friday, matching put and call contracts with identical contract specifications for trades occurring within 5 minutes from each other. Then, in order to move from intraday (intra-Friday) data to weekly time series, and given that such an option equivalent is constructed for the first time, various approximations are adopted for the representative weekly quote. First, we employ the last trade of each Friday (*Last Trade*), to conform to the forward practices in the literature. Second, we construct the Friday's average, that is the mean of the distribution of the intraday synthetic forward on each Friday (*Average*). A third measure is constructed from the mean of at-the-money contracts (*ATM*), thus screening what are typically the most frequently traded contracts. Fourth, we consider the median of the distribution of the intraday synthetic forward (*Median*). Lastly, we employ the volume of trade for each contract as a weight and calculated the weighted Friday's average of synthetic forward contracts (*W. Average*). Note that the above techniques are applied to construct both the synthetic forward and its respective spot rate, in order to match the option equivalent as closely as possible with the corresponding spot exchange rate on each Friday.

This seemingly simple process is confronted with several challenges. A drawback in our data set is the decline in trading after 1995, which became apparent around 1997 and onwards. Although a straightforward explanation for such decline in trading is not provided by the PHLX, possible explanations gathered through our telephone interviews with PHLX managers include the gradual shift of focus to electronic trading

[14] In cases of no data on Fridays (e.g. due to public holidays) we use the immediately preceding day within the same week.

and, most predominantly, the investors' increasing preference for the United Currency Options Market (UCOM), a November 1994 innovation of the PHLX, which offers the possibility to customise the contract specifications[15]. The above drop in observations weighted on our effort to construct weekly estimates from a comparable number of intraday observations. For robustness, we therefore also create a different set of series where we removed outliers, defined as observations of the option equivalent in the 5th and 95th quantiles of the distribution. We use the outlier-removed sample to check the robustness of our empirical results; as discussed in our empirical work below, our results are qualitatively identical for these two different sets of data.

Another challenge involves matching the conventions of the standardised (options exchange) market with the customised (OTC) forward market in terms of maturities and expiry dates. In the forward market, the expiry date of the contract can be on any day of the month, for contracts of any conventional maturities used in the forward market. The time to maturity becomes an immediately observable feature of forward contracts. On the contrary, the PHLX offers only two expiry days per month–namely mid-month and month-end expiries respectively. Therefore, the observable features of the trades on the PHLX are the expiry date and, consequently, the type of contract–recall that we only use mid-month contracts in this chapter. This implies that for the forward market we can observe the maturity of the contract expiring on each day, whereas in the PHLX system options contracts of all maturities expire on a specific day of each month[16]. Naturally, these differences reduce the comparability of forward and option equivalent rates[17].

[15]Up to that point, the PHLX only offered standardised contract specifications and market conventions, i.e. contracts that specify the currency pair traded, the contract size, strike price intervals, expiration dates, price quoting and premium settlement. UCOM increased flexibility by introducing customised currency options. This offered a choice to investors over all aspects of a currency option trade (exercise price, selection of currency pairs, premium quotation as either units of currency or percent of underlying value, and customised expiration dates of up to two years).

[16]The PHLX has standardised expiry dates, by setting specific expiry conventions. Mid-month contracts expire only on the first Friday following the third Wednesday of the expiry month, and month-end contracts expire on the last Friday of the month. The contracts trade on a fixed-months quarterly cycle (March, June, September, December and the two months following the current month).

[17]See Breuer and Wohar (1996) for a discussion of specific institutional characteristics of forward

For the construction of the 1- and 3-month synthetic forward contract we fix the expiry date and gather all the relevant contracts that have been traded 1 month (4 weeks) and 3 months (12 weeks) ahead. Every time the expiry date is reached, the contracts of the expiry date enter the new cycle with a next expiry date. Therefore, at a specific time all contracts selected will expire on the same specific date, although the contracts might have begun trading at different points in the past. As a result, our 1- and 3-month synthetic forward contract is different from our 1- and 3-month conventional forward contract in that the former has a specific expiry date, based on the expiry cycle of the options, as specified by the PHLX, and includes all contract maturities trading within these dates, whereas the latter has fixed time to maturity but can expire any day of the month. This feature of the synthetic forward allows us the flexibility of assigning different values for k in the Fama regression given by equation (2.2) when we use the option equivalent. Namely, k can take the values 4, 8, 12, 16, 24, 36, 52 (weeks), in contrast with the forward rate which in the literature is typically used for 4, 8 or 12 (weeks). In this chapter, however, we study the 4-, 8- and 12-week contracts for our synthetic forward contracts to make a tighter comparison with the relevant literature.

The resulting series of interest are as follows. For tests based on forward contracts, the data set includes the logarithm of the spot exchange rate, s_t and the logarithm of the 1- and 3-month forward exchange rates, f_t^4 and f_t^{12} respectively, at weekly frequency–specifically, end-of-the-week prices. For tests based on options contracts, the series of interest consist of the logarithm of the five different definitions for the synthetic forward at time t, (o_{t,j_h}^k) and the respective spot $(s_{t,h})$, where the subscript $j =$ *ATM, Average, Last Trade, Median* and *W. Average* corresponds to the five different methods of data construction described above; the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively; and $k = 4, 8, 12$ is the maturity of the contract in weeks. This data set provides a variety

markets that are relevant to tests of unbiasedness.

of distinct sets of spot and synthetic forward rates for each of the three exchange rates we examine. Given the vast amount of results we obtained, the core of the empirical work is based on weekly data for s_t and f_t^4 for forward-based tests, and $s_{t,h}$ and o_{t,j_h}^4 for options-based tests, while we shall use the remaining data in our robustness analysis.

2.4 Empirical results

2.4.1 Preliminary data analysis

Several unit root tests are conducted to shed light on the integration properties of the time series under investigation–the tests include the Phillips-Perron test and the more recent MZ_α and MZ_t tests proposed by Ng and Perron (2001) for the null hypothesis of a unit root[18]. In keeping with the large number of studies of unit root behaviour for FX time series, we are in each case unable to reject the unit root null hypothesis for s, f and o, at conventional nominal levels of significance for the level series. On the other hand, differencing the series appears to induce stationarity in each case. Hence, the unit root tests clearly indicate that spot, forward and option equivalent rates time series are realisation from stochastic processes integrated of order one[19].

While unit root tests on s, f and o provide results that are consistent with conventional wisdom in the profession, a more controversial issue is whether the forward premium is a unit root process, whereas the integration properties of the option premium have–to the best of our knowledge–not been studied to date. We apply unit root tests to the forward premium, $\left(f_t^k - s_t\right)$ and the option premium, $\left(o_{t,h}^k - s_{t,h}^k\right)$. The results, reported in Table 2.1, indicate that, for each forward premium and each option premium examined, the null hypothesis of a unit root is rejected at conventional significance levels. In turn, this result suggests that the two variables in the Fama regression

[18]The Ng-Perron tests use generalised least squares-detrending to maximise test power and a modified information criterion to select the lag truncation in order to minimise size distortion (see Ng and Perron, 2001).

[19]These test statistics are not reported to conserve space but they are available from the authors upon request.

given in equation (2.2), namely the exchange rate change and the premium, are both mean reverting. This is an important preliminary finding, given the well-documented difficulties in detecting mean reversion in the forward premium due to its long memory properties that make the forward premium behave like a fractionally integrated process (Baillie and Bollerslev, 1994, 2004).

2.4.2 Fama regressions

Our next exercise is to estimate the conventional Fama regression, equation (2.2) for each currency pair and type of forward rate and option equivalent. This would in principle show the existence of a "forward bias," as recorded in much previous research, and address the question whether there is a similar "option bias" when estimating the Fama regression with the option equivalent.

The results, reported in Table 2.2, are consistent with the existence of both forward and option bias. Panel A presents the estimation results for the conventional forward contract. We observe that the constant term α is close to zero and often statistically insignificant, whereas β, albeit positive except for the case of the yen, is always estimated to be statistically insignificant. The results are somewhat similar for the option equivalent (Panel B of Table 2.2). The constant terms are, in most cases, small and insignificantly different from zero. The best estimate of β is for the Swiss franc, where we find positive and significant estimates of β, but the magnitude is close to zero. For all other cases, the slope coefficient β is statistically insignificantly different from zero.

Overall, the results in Table 2.2 suggest that estimation of the Fama regression using our option equivalent measure rejects unbiasedness and indicates the existence of an option bias puzzle (Panel B) that is consistent with the stylised facts leading to the forward bias puzzle[20] (Panel A).

[20] Asymptotic standard errors were calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals throughout the paper (Newey and West, 1987).

2.4.3 Cointegration tests

As discussed in Section 2.2.1, the Fama regression (2.2) may not be appropriate for testing the unbiasedness hypothesis because endogeneity issues and an omitted risk premium may render these tests uninformative (e.g. Barnhart, McNown and Wallace, 1999) and because the potentially unbalanced nature of the Fama regression makes statistical inference cumbersome in that regression (e.g. Baillie and Bollerslev, 1994, 2004). Regression (2.2) is essentially a stringent test of UIP under the risk-neutral rational-expectations FX market efficiency hypothesis, which is a sufficient but not necessary condition for unbiasedness.

In this section we shift our attention to cointegration analysis, by applying several cointegration tests. We begin with the test proposed by Phillips and Loretan (1991), based on a nonlinear least squares (NLS) estimation procedure which accounts for endogeneity of the regressors. This test is particularly straightforward to implement in that it has a known standard asymptotic distribution, allowing us to test the unbiasedness hypothesis as a test that s_{t+k} and f_t^k cointegrate for the case of forward contracts and that s_{t+k} and o_t^k cointegrate for the case of options contracts–for the case of options, the tests are conducted for all the various definitions of the option equivalent o_t^k given in Section 2.3.2. The formal test of unbiasedness involves testing cointegration and the hypothesis that $\beta = 1$, where now β is a cointegrating parameter, and then testing the hypothesis that the residuals from the cointegration test are white noise.

Our results, reported in Table 2.3, show ample support in favour of cointegration between the future spot rate and the current forward rate or the current synthetic forward (option equivalent) rate. Panel A and B display the results for the conventional forward and the option equivalent respectively. The results are very similar. The slope coefficient β (which is now a cointegrating parameter) is generally very close to unity for all cases. The formal test that the cointegrating relationship is of the form $[1, -1]$ is generally not rejected. Also, the statistically significant values of the ADF tests on the residuals from the auxiliary regressions support the hypothesis of cointegration (i.e. reject the null hypothesis of no cointegration).

In addition to the Phillips-Loretan tests, we also carry out a more general nonparametric cointegration test, introduced by Bierens (1997a). This test is appealing in the present context and is deemed superior to standard parametric tests since it has been shown to be capable of detecting cointegration when the data generating process is nonlinear[21]. Although it follows the spirit of reduced rank cointegration tests, the Bierens methodology can consistently estimate the number of cointegrating vectors and also test for parametric restrictions on the cointegrating vectors, on the basis of the ordered solutions of a generalised eigenvalue problem. As for verifying the numbers of cointegrating vectors, Bierens shows how to calculate a λ_{min} test, which is analogous to the Johansen trace test, by testing the null of lower against higher numbers of cointegrating vectors. However, Bierens (1997b) considers the λ_{min} test a tentative outcome and suggests a double check on it by presenting a method for estimating the number of cointegrating vectors ($g_m(r_o)$ test).

The results from performing the Bierens test are reported in Table 2.4 (Panels A and B for the conventional forward and the option equivalent respectively). These results, which are again similar between the two different derivatives examined, again provide empirical evidence in favour of cointegration between the spot and the lagged (synthetic) forward rate. Notably, the results from the λ_{min} tests and the $g_m(r_o)$ indicate always cointegration at the 5 percent significance level and suggest the existence of a unique cointegrating vector. Further tests that specify the form of the vector spanning the cointegration space by imposing the restriction of $[1, -1]$ indicated that the null hypothesis of a one-to-one cointegrating relationship could not be rejected at conventional significance levels by the relevant trace test.

Finally, we employ the Dynamic OLS (DOLS) estimation method to provide further

[21] Presence of nonlinearity in spot-forward models has been argued by several authors; e.g. see Engel and Hamilton (1990), Clarida, Sarno, Taylor and Valente (2003), Sarno, Valente and Leon (2005).

evidence on the existence of a one-to-one cointegrating relationship between the spot and the lagged (synthetic) forward rate. While we know the asymptotic distribution of the two previous tests employed, we know little about their small sample properties. However, Stock and Watson (1993) document that DOLS has better small sample characteristics than, for example, the Phillips-Loretan test for the purpose of estimating efficiently the cointegrating vector. The results, reported in Table 2.5, provide further comforting evidence that the cointegrating parameter is unity for each of the currencies and derivatives instruments examined.

Overall, all cointegration techniques employed yield the same outcome, providing ample support in favour of a one-to-one cointegrating relationship between the spot and the lagged (synthetic) forward rate. This is an encouraging result given the difficulties that a large empirical literature finds in detecting an exactly proportional cointegrating relationship between spot and forward rates (e.g. Maynard, 2003). However, this is a necessary condition towards establishing FX unbiasedness, albeit not yet sufficient.

2.4.4 Residuals tests for FX forward unbiasedness

Following Barnhart, McNown and Wallace (1999), as an additional and more stringent test for unbiasedness, we move on to examine the residual correlation of the errors arising from the cointegrating relationship between the spot and the lagged (synthetic) forward. Given the earlier empirical evidence on the existence of a $[1, -1]$ cointegrating vector in the relationship between the spot and the (synthetic) forward, we construct the deviation from UIP as the difference of the lagged (synthetic) forward rate from the spot rate–i.e. we impose the $[1, -1]$ cointegrating vector–thus generating "restricted" cointegrating residuals. We then employ tests of residual serial correlation, regressing the residuals of each series on their own lagged values (including 4 lags for the 4-lag options and forward series), and a test of cross-correlation, where the residuals are regressed on their own lagged values and on lagged values of the other series (employing again four lags for each different series). We then perform a joint significance coefficient restriction (Wald) test for the null hypothesis of no residual correlation, against the alternative that at least one lag is statistically significant.

The results, presented in Table 2.6 (Panels A and B), are again very similar between the options and the forward case. Indeed, for all three currencies and for both conventional and synthetic forward rates, the relevant F-test cannot reject the null of no residual autocorrelation and no cross correlation at conventional significance levels (with the only exception of the Swiss franc in one case). Overall, the outcome points towards the validity of the unbiasedness hypothesis[22].

2.5 Robustness analysis

In this section we discuss several robustness checks carried out in order to evaluate the sensitivity of the empirical results reported in the previous section. In particular, we assess the robustness of our results: (a) to the choice of the number of lags employed in the synthetic forward for the case of 1-month contracts, and (b) to the choice of the maturity of the (synthetic) forward contract, switching to 3-month contracts. The results relating to this section are not reported to conserve space, but they are available upon request.

2.5.1 1-month contracts

Given that our synthetic forward contract contains a mixture of different maturities for the same expiry date, we experiment with taking different numbers of lags, corresponding to different maturities. For that we select 8 lags, corresponding to a maturity of 2 months, for which we run the same regressions considered in the core analysis. Our analysis focuses on the synthetic forward for each currency, on both the original sample

[22] We also performed the same tests on the Phillips-Loretan residuals (not reported but available upon request). These results are qualitatively similar to the results reported for the "restricted" residuals.

and the void of outliers sample.

The results confirm the similarities between the conventional and the synthetic forward and show no qualitative difference from the case with 4 lags. Namely, unbiasedness is again rejected on the basis of the Fama regression, with the estimates of both the constant α and the slope β being virtually the same as in the core analysis, confirming the existence of an option bias.

On the contrary, ample evidence of a cointegrating relationship between the spot and the synthetic forward is suggested by cointegration tests, which indicate the presence of a $[1, -1]$ cointegrating vector for all cases examined.

Lastly, we perform an autocorrelation test on the "restricted" residuals, generated with the same method as in the core analysis; however, this time, 8 lags are employed for each currency in the own- and cross-correlation tests. The majority of outcomes cannot reject the null of no autocorrelation at conventional significance levels. Nevertheless, there are minor exceptions where serial or cross correlation is detected, but they do not entice any specific pattern. Thus, this evidence notwithstanding, we conclude that our core results are robust to changes in the lags employed and offer support to the unbiasedness hypothesis on the basis of cointegration and residuals tests.

2.5.2 3-month contracts

We then re-estimate the core regressions for each exchange rate examined using a 3-month forward contract and a 3-month synthetic forward contract, at the weekly frequency, to assess the robustness to the choice of the contract maturity. In order to construct our synthetic contract we choose the expiry dates of the fixed quarterly cycles (March, June, September and December), and gathered all relevant contracts. Again, we have 3-month synthetic forwards with a range of maturities from 4 to 52 weeks. For reasons of consistency to the forward case, we choose to work with 12 lags.

The results are, again, very similar between the forward and the synthetic forward case and between the 1-month and the 3-month contracts. The Fama regression for the conventional forward contract presents negative and insignificant coefficients for the β slope parameter, whereas the constant term α is estimated to be close to zero (albeit significant for the yen and the Swiss franc). These results are comparable with the ones obtained from estimating the Fama regression with the option premium (equation (2.7)).

Shifting our attention to the cointegration tests, all three cointegration tests used in this chapter detect the existence of a $[1, -1]$ cointegrating relationship in the case of the spot-forward as well as spot-option cases. Finally, the autocorrelation and cross-correlation tests on the residuals (this time performed with 12 lags) lead to similar conclusions as in the core analysis–i.e. the null of no autocorrelation cannot be rejected.

2.6 Conclusions

Armed with several tests proposed by the literature testing forward rate unbiasedness in the FX market, this chapter provides a simple, yet intuitive bridge to a different derivatives market, the currency options market, as a vehicle of forming expectations about future spot exchange rates. Our main focus is on performing tests of the unbiasedness hypothesis. To that end, we used the conventional forward rate and also introduced an option equivalent (synthetic forward) contract. We then apply some prominent tests of unbiasedness, based on the standard UIP condition in a stationary setting as well as cointegration tests for unbiasedness of the (synthetic) forward rate, the latter combined with residual autocorrelation tests.

This research provides encouraging results. We manage to bridge the distance between the forward (OTC) market and the options (exchange traded) market, by directly comparing the test results obtained for the two markets. Viewed from a different angle, our research offers a novel robustness check to the tenacity of the well-documented forward bias anomaly that characterises the relevant literature.

We record no qualitative difference between the two types of derivatives products in our results. Specifically, our results suggest the existence of an "options bias," similar to the forward bias, frequently recorded in the relevant literature estimating stationary regressions of the exchange rate change on the lagged forward premium. This finding indicates, in turn, violation of market efficiency in its risk neutral formulation as implied by UIP, possibly as a consequence of the existence of a risk premium or possibly because of the estimation problems induced by the persistence of the forward (option) premium and of exchange rate volatility. We therefore shift our attention to a cointegration and a residual correlation analysis that allows for the endogeneity problems caused by a potential unobserved risk premium term and requires less stringent conditions on the error term process. This analysis attests that indeed the (synthetic) forward is an unbiased predictor of the future spot rate.

Overall, we interpret the evidence in this chapter as suggesting that forward and options provide optimal exchange rate predictions consistent with the notion of unbiasedness.

Table 2.1 Unit root tests on the (synthetic) forward premium: 1-month contracts

Panel A) Forward contracts

	PP	MZ_a	MZ_t
GBP	-3.646	-15.35	-3.096
JPY	-3.649	-14.92	-3.625
CHF	-3.894	-17.11	-4.173

Panel B) Option contracts

		original sample					outliers removed		
		PP	MZ_a	MZ_t			PP	MZ_a	MZ_t
GBP	ATM	-3.485	-14.64	-4.302	GBP	ATM	-5.926	-15.18	-3.185
	Average	-4.011	-14.81	-5.220		Average	-4.674	-15.99	-4.942
	Last Trade	-4.725	-15.64	-4.081		Last Trade	-4.684	-15.74	-4.038
	Median	-3.199	-14.33	-4.596		Median	-4.993	-14.86	-5.425
	W. average	-3.494	-14.28	-4.980		W. average	-4.030	-12.40	-4.476
JPY	ATM	-5.864	-15.11	-5.884	JPY	ATM	-5.924	-11.41	-4.409
	Average	-5.522	-15.45	-5.931		Average	-5.796	-12.30	-4.325
	Last Trade	-4.514	-13.75	-5.385		Last Trade	-4.528	-12.10	-4.825
	Median	-3.527	-16.13	-4.662		Median	-4.903	-13.19	-3.569
	W. average	-3.111	-15.36	-4.940		W. average	-4.133	-13.68	-3.748
CHF	ATM	-4.044	-14.14	-4.229	CHF	ATM	-4.398	-15.79	-3.227
	Average	-3.266	-14.12	-4.148		Average	-3.544	-14.66	-3.781
	Last Trade	-4.304	-14.86	-4.915		Last Trade	-4.696	-16.46	-3.028
	Median	-5.048	-14.83	-3.521		Median	-4.576	-11.23	-4.311
	W. average	-3.177	-12.95	-2.986		W. average	-3.613	-15.71	-3.961

Notes. *Panel A)* of the table presents unit root test statistics on the forward premium, $(f_t^k - s_t)$, whereas *Panel B)* presents unit root test statistics on the synthetic forward premium, $(o_{t,h}^k - s_{t,h}^k)$, where j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 2.3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. For both panels, the test statistics are tests of the null hypothesis of a unit root. The asymptotic critical values, at the 1% and 5% significance level respectively, for the Phillips-Perron (PP) test are -2.567 and -1.941; for the Ng and Perron (2001) MZ_a test they are -13.800 and -8.100; and for the Ng and Perron (2001) MZ_t test they are -2.580 and -1.980.

Table 2.2 Fama regressions: 1-month contracts (k=4)

Panel A) Forward premium Fama regressions

	α	SE(α)	β	SE(β)
GBP	0.002*	(0.001)	0.371	(0.292)
JPY	0.003**	(0.001)	-0.260	(0.248)
CHF	0.002**	(0.001)	0.148	(0.274)

Panel B) Option premium Fama regressions

		original sample				outliers removed			
		α	SE(α)	β	SE(β)	α	SE(α)	β	SE(β)
GBP	ATM	0.001	(0.001)	0.060	(0.065)	0.001	(0.001)	0.088	(0.070)
	Average	0.001	(0.001)	0.021	(0.065)	0.001	(0.001)	0.027	(0.078)
	Last Trade	0.001	(0.001)	-0.039	(0.050)	0.001	(0.001)	-0.040	(0.060)
	Median	0.001	(0.001)	0.001	(0.057)	0.001	(0.001)	0.024	(0.067)
	W. average	0.001	(0.000)	-0.011	(0.056)	0.001	(0.001)	-0.033	(0.069)
JPY	ATM	0.003**	(0.001)	-0.009	(0.047)	0.002**	(0.001)	-0.126	(0.110)
	Average	0.003**	(0.001)	-0.031	(0.047)	0.002**	(0.001)	-0.197*	(0.080)
	Last Trade	0.003**	(0.001)	-0.023	(0.034)	0.003**	(0.001)	-0.082	(0.052)
	Median	0.003**	(0.001)	-0.016	(0.040)	0.003**	(0.001)	-0.094	(0.064)
	W. average	0.003**	(0.001)	-0.028	(0.042)	0.002**	(0.001)	-0.107***	(0.064)
CHF	ATM	0.002	(0.001)	0.135*	(0.065)	0.002	(0.001)	0.165**	(0.075)
	Average	0.002	(0.001)	0.117	(0.064)	0.002	(0.001)	0.262*	(0.080)
	Last Trade	0.002	(0.001)	0.060	(0.049)	0.002	(0.001)	0.132*	(0.052)
	Median	0.002	(0.001)	0.089	(0.050)	0.002	(0.001)	0.063*	(3.420)
	W. average	0.002*	(0.001)	0.044	(0.052)	0.002	(0.001)	0.068**	(2.045)

Notes. *Panel A)* The table shows the results from estimating, by ordinary least squares, the conventional forward premium (Fama) regression in equation (2.2). *Panel B)* The table shows the results from estimating, by ordinary least squares, the option premium (Fama) regression in equation (2.7). For both panels, figures in parentheses (SE(α) and SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals up to the third decimal point (Newey and West, 1987). One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Table 2.3 Phillips-Loretan cointegration tests: 1-month contracts (k=4)

Panel A) Spot-forward relationship

	β	ADF	F
GBP	1.003	-15.268	[0.236]
JPY	0.993	-15.908	[0.318]
CHF	1.011	-16.031	[0.792]

Panel B) Spot-option relationship

		original sample			outliers removed		
		β	ADF	F	β	ADF	F
GBP	ATM	1.006	-15.562	[0.563]	1.007	-15.585	[0.222]
	Average	1.025	-15.739	[0.531]	1.021	-15.702	[0.080]
	Last Trade	0.978	-15.759	[0.720]	1.123	-15.593	[0.002]
	Median	1.125	-15.773	[0.793]	1.025	-15.853	[0.046]
	W. average	0.972	-15.979	[0.782]	0.988	-15.890	[0.097]
JPY	ATM	0.988	-15.200	[0.660]	0.997	-15.713	[0.391]
	Average	0.991	-15.117	[0.650]	0.998	-15.584	[0.574]
	Last Trade	0.990	-15.064	[0.653]	0.996	-15.409	[0.445]
	Median	0.987	-14.857	[0.745]	0.995	-15.434	[0.805]
	W. average	0.989	-14.890	[0.691]	0.996	-15.278	[0.380]
CHF	ATM	1.008	-15.466	[0.650]	1.008	-15.608	[0.533]
	Average	1.003	-15.491	[0.908]	1.003	-15.650	[0.369]
	Last Trade	1.010	-15.516	[0.829]	1.010	-15.445	[0.421]
	Median	1.003	-15.489	[0.919]	1.004	-15.874	[0.270]
	W. Average	1.012	-15.585	[0.723]	1.012	-15.666	[0.430]

Notes. *Panel A)* The table presents the results from testing for cointegration between the spot rate, s_{t+4} and the forward rate, f_t^4 using the Phillips-Loretan (1991) test. *Panel B)* The table presents the results from testing for cointegration between the spot rate, s_{t+4} and the synthetic forward, o_{t,j_h}^4 rate using the Phillips-Loretan (1991) test, where j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 2.3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. *For both panels*, β denotes the cointegrating parameter. ADF is the Augmented Dickey Fuller test statistic for a unit root in the residuals (i.e. for no cointegration). The column F gives the p-value from the relevant F-statistic for the null hypothesis that the cointegrating vector is $[1, -1]$.

Table 2.4 Bierens nonparametric cointegration tests: 1-month contracts (k=4)

Panel A) Spot-forward relationship

	$\lambda_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
GBP	[0.000]*	7.30×10^7	1.020
	[0.078]	$\mathbf{1.96 \times 10^0}$	
		1.02×10^4	
JPY	[0.000]*	3.68×10^5	1.090
	[0.423]	$\mathbf{1.31 \times 10^1}$	
		2.02×10^6	
CHF	[0.000]*	5.61×10^4	1.370
	[0.137]	$\mathbf{8.18 \times 10^2}$	
		1.33×10^7	

(continued ...)

(... Table 2.4 continued)

Panel B) Spot-option relationship

GBP)

	original sample			outliers removed		
	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
ATM	[0.000]*	3.21×10^{1}	1.010	[0.000]*	3.89×10^{0}	1.010
	[0.062]	$\mathbf{7.20 \times 10^{-3}}$		[0.062]	$\mathbf{5.88 \times 10^{-3}}$	
		2.39×10^{1}			1.97×10^{1}	
Average	[0.000]*	7.17×10^{9}	1.140	[0.000]*	8.74×10^{13}	1.050
	[0.058]	$\mathbf{3.62 \times 10^{-2}}$		[0.061]	$\mathbf{2.72 \times 10^{-6}}$	
		1.09×10^{2}			8.79×10^{-3}	
Last Trade	[0.000]*	4.38×10^{1}	1.010	[0.000]*	8.36×10^{11}	1.010
	[0.064]	$\mathbf{4.90 \times 10^{-3}}$		[0.063]	$\mathbf{2.68 \times 10^{-4}}$	
		1.75×10^{1}			9.18×10^{-1}	
Median	[0.000]*	2.31×10^{12}	1.080	[0.000]*	1.37×10^{12}	1.050
	[0.059]	$\mathbf{1.08 \times 10^{-4}}$		0.059	$\mathbf{1.85 \times 10^{-4}}$	
		3.33×10^{-1}			5.59×10^{-1}	
W. Average	[0.000]*	1.34×10^{12}	1.040	[0.000]*	5.43×10^{11}	1.020
	[0.061]	$\mathbf{1.77 \times 10^{-4}}$		[0.061]	$\mathbf{4.36 \times 10^{-4}}$	
		5.73×10^{-1}			1.41×10^{0}	

JPY)

	original sample			outliers removed		
	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
ATM	[0.000]*	1.07×10^{11}	1.040	[0.000]*	2.08×10^{7}	1.100
	[0.370]	$\mathbf{5.94 \times 10^{-5}}$	1.040	[0.377]	$\mathbf{2.93 \times 10^{-1}}$	
		7.05×10^{0}			3.62×10^{4}	
Average	[0.000]*	3.28×10^{10}	1.180	[0.000]*	6.47×10^{7}	1.190
	[0.368]	$\mathbf{1.96 \times 10^{-4}}$	1.180	[0.380]	$\mathbf{9.50 \times 10^{-2}}$	
		2.30×10^{1}			1.17×10^{4}	
Last Trade	[0.000]*	3.77×10^{9}	1.020	[0.000]*	1.88×10^{7}	1.180
	[0.378]	$\mathbf{1.61 \times 10^{-3}}$	1.020	[0.379]	$\mathbf{3.21 \times 10^{-1}}$	
		2.00×10^{2}			4.02×10^{4}	
Median	[0.000]*	6.11×10^{8}	1.550	[0.000]*	1.99×10^{8}	1.270
	0.362	$\mathbf{1.08 \times 10^{-02}}$	1.550	0.370	$\mathbf{3.19 \times 10^{-2}}$	
		1.23×10^{3}			3.80×10^{3}	
W. Average	[0.000]*	1.96×10^{10}	1.190	[0.000]*	7.12×10^{7}	1.290
	[0.364]	$\mathbf{3.33 \times 10^{-04}}$	1.190	[0.371]	$\mathbf{8.88 \times 10^{-2}}$	
		$3.84\mathrm{E} \times 10^{1}$			1.06×10^{4}	

(continued ...)

(... Table 2.4 continued)

CHF)

	original sample			outliers removed		
	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
ATM	[0.000]*	2.55×10^7	1.180	[0.000]*	9.92×10^6	1.210
	[0.120]	$\mathbf{2.39 \times 10^0}$		[0.121]	$\mathbf{6.03 \times 10^0}$	
		3.01×10^4			7.73×10^4	
Average	[0.000]*	5.49×10^7	1.150	[0.000]*	1.29×10^7	1.170
	[0.120]	$\mathbf{1.12 \times 10^0}$		[0.118]	$\mathbf{4.85 \times 10^0}$	
		1.40×10^4			5.97×10^4	
Last Trade	[0.000]*	5.38×10^7	1.160	[0.000]*	1.92×10^7	1.170
	[0.123]	$\mathbf{1.07 \times 10^0}$		[0.121]	$\mathbf{3.09 \times 10^0}$	
		1.43×10^4			3.99×10^4	
Median	[0.000]*	3.05×10^7	1.170	[0.000]*	8.68×10^6	1.170
	[0.120]	$\mathbf{1.99 \times 10^0}$		[0.118]	$\mathbf{7.28 \times 10^0}$	
		2.51×10^4			8.84×10^4	
W. Average	[0.000]*	1.06×10^8	1.130	[0.000]*	2.08×10^7	1.150
	[0.119]	$\mathbf{5.79 \times 10^1}$		[0.118]	$\mathbf{3.04 \times 10^0}$	
		7.22×10^3			3.69×10^4	

Notes. The tables present the results from the nonparametric cointegration tests of Bierens (1997a) applied to the spot-forward relationship (s_{t+4} and f_t^4 - *Panel A*) and the spot-option relationship (s_{t+4} and o^4_{t,j_h} - *Panel B*); j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 2.3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. The first column of results ($\lambda_{\min}$) shows the p-values of the $\lambda_{\min}$ test statistic for T1 (which is H_o: $r=0$ *vs.* H_1: $r=1$) and for T2 (which is H_o: $r=1$ *vs.* H_1 $r=2$) respectively. The second column calculates $r_m = \arg\min_{r_o \le 2}\{g_m(r_o)\}$ for $m = 2$, where r_o is the number of cointegrating vectors; the table presents the r_m values for $r_o = 0, 1, 2$. The number in bold emphasizes the minimum r_m value, which indicates the number of cointegrating relationships identified by the Bierens test. The final column presents the results from the trace test for the null hypothesis that the cointegrating vector is $[1, -1]$, i.e. $H_o : \beta' = (1, -1)$; the appropriate 5-percent critical value is 4.70. The asterisk denotes statistical significance at the 5 percent level.

Table 2.5 DOLS cointegration relationships: 1-month contracts (k=4)

Panel A) Spot-forward relationship

	β	F
GBP	1.004	[0.151]
JPY	1.000	[0.640]
CHF	0.995	[0.286]

Panel B) Spot-option relationship

		original sample		outliers removed	
		β	F	β	F
GBP	ATM	1.001	[0.722]	1.001	[0.606]
	Average	1.003	[0.412]	1.003	[0.356]
	Last Trade	1.003	[0.379]	1.004	[0.292]
	Median	1.001	[0.790]	1.001	[0.685]
	W. average	1.003	[0.479]	1.003	[0.449]
JPY	ATM	0.998	[0.117]	0.999	[0.146]
	Average	0.998	[0.107]	0.999	[0.166]
	Last Trade	0.998	[0.138]	0.999	[0.140]
	Median	0.999	[0.121]	0.999	[0.228]
	W. average	0.998	[0.111]	0.999	[0.032]
CHF	ATM	0.998	[0.761]	0.996	[0.374]
	Average	0.999	[0.983]	0.997	[0.597]
	Last Trade	0.998	[0.806]	0.997	[0.657]
	Median	1.002	[0.664]	0.999	[0.841]
	W. Average	0.998	[0.779]	0.996	[0.420]

Notes. *Panel A)* The table presents the estimates of the cointegrating relationship between the spot rate, s_{t+4} and the forward rate, f_t^4 using the Dynamic-OLS (DOLS) method (1988). *Panel B)* The table presents the estimates from the cointegrating relationship between the spot rate, s_{t+4} and the synthetic forward, o_{t,j_h}^4 rate using the Dynamic OLS (DOLS) method; j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 2.3.2, and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. *For both panels*, β denotes the cointegrating parameter. The column F gives the p-value from the relevant F-statistic for the null hypothesis that the cointegrating vector is $[1, -1]$.

Table 2.6 Residual tests: 1-month contracts (k=4)

Panel A) Residual correlation tests for spot-forward

	AC	CC
GBP	[0.399]	[0.862]
JPY	[0.327]	[0.322]
CHF	[0.178]	[0.543]

Panel B) Residual correlation tests for spot-options

		original sample		outliers removed	
		AC	CC	AC	CC
GBP	ATM	[0.540]	[0.836]	[0.610]	[0.957]
	Average	[0.602]	[0.790]	[0.764]	[0.831]
	Last Trade	[0.755]	[0.827]	[0.445]	[0.852]
	Median	[0.534]	[0.793]	[0.663]	[0.903]
	W. average	[0.530]	[0.769]	[0.787]	[0.855]
JPY	ATM	[0.087]	[0.369]	[0.063]	[0.128]
	Average	[0.119]	[0.165]	[0.079]	[0.279]
	Last Trade	[0.179]	[0.073]	[0.318]	[0.187]
	Median	[0.143]	[0.082]	[0.271]	[0.151]
	W. average	[0.148]	[0.126]	[0.055]	[0.166]
CHF	ATM	[0.194]	[0.265]	[0.436]	[0.688]
	Average	[0.066]	[0.132]	[0.352]	[0.570]
	Last Trade	[0.167]	[0.285]	[0.447]	[0.485]
	Median	[0.062]	[0.072]	[0.148]	[0.177]
	W. Average	[0.019]*	[0.026]*	[0.247]	[0.179]

Notes. The tables present the p-values for the relevant F-statistics for joint coefficient restriction on $s_{t+k} - f_t$ for the forward (*Panel A*) and $s_{t+k} - o_{t,j}$ for the synthetic forward (*Panel B*), where j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 2.3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. We perform tests for autocorrelation (AC) and cross correlation (CC) for the spot-forward and the spot-option relationships. For the case of the forward, the F-statistic is distributed with (4, 925) degrees of freedom for the AC test and (12, 917) degrees of freedom for the CC test respectively. For the case of options the F-statistic is distributed with (4, 922) degrees of freedom for the AC test and (12, 914) degrees of freedom for the CC test respectively. The asterisk denotes statistical significance at the 5 percent level.

Chapter Three

3 The behaviour of the real exchange rate: Evidence from regression quantiles

3.1 Introduction

3.1.1 Real exchange rate issues and related literature

Purchasing Power Parity (PPP) has long been considered as one of the fundamental arbitrage laws in international asset pricing. The building block of PPP is the Law of One Price (LOP), which contends that, in the absence of arbitrage, identical goods should be selling at the same price across countries. Aggregating across all tradable goods in an economy, we obtain PPP, which suggests that price levels between two countries should be equal, if expressed in the same currency. PPP, therefore, provides an equilibrium relationship for the real exchange rate (RER), which is the nominal exchange rate, adjusted for relative price levels. If PPP holds, the relative price levels and/or the bilateral nominal exchange rate would adjust in such a way so that the RER remain constant. In that sense, variations in the RER would suggest deviations from PPP.

Although intuitive theoretically, in practice the RER exhibits high variability over time and spends long periods away from its suggested PPP equilibrium. The ambiguity surrounding the persistency of the RER and the validity of PPP, is well summarised into two relevant puzzles. The first one directly investigates the persistency of the RER process. As long as the RER is reverting back to its PPP equilibrium, albeit slowly, this implies that PPP should, at least, be seen as a long term anchor for determining the RER equilibrium value, although it may not be holding at each point in time. However, if the deviations from the PPP are permanent, this suggests the absence of a unique, constant equilibrium. The second puzzle (Rogoff, 1996) is trying to rationalise the persistence of the RER and reconcile its extremely volatile nature in the short run with the extremely slow rate at which shocks appear to damp out. This puzzle raises the issue of the types and role of the shocks that hit the RER and how they impact on the RER mean reversion[23],[24].

Given the importance of PPP in international finance and our limited understanding of the RER behaviour, an extensive amount of research is being dedicated to testing the unit root hypothesis in the RER. Evidence from early attempts was clearly rejecting PPP (for a summary exposition of early tests see Sarno and Taylor, 2003). Nevertheless, it soon became obvious that standard unit root tests have low power in rejecting the null of a unit root. This shortcoming is nurtured by the inability of these tests to take into account certain distributional stylised facts of the exchange rates in general, and the RER in particular[25]. More precisely, although the true RER distribution of the RER is not known, the notion that it is normally distributed is refuted, because the overall process appears to be better described by leptocurtic distributions (McLachlan and Peel, 2000). The non-normality of the RER distribution confounds standard unit root tests, by lowering their power (Perron, 1990; Kim, Nelson and Startz, 1998).

Parametric unit root tests of increased sophisticated and complex structures, which accounted for the non-normality of the RER, offered more robust alternatives (Pippenger and Goering, 1993, 2000; Michael, Nobay and Peel, 1997; Nelson, Piger and Zivot, 2001). These tests would typically result from regime switching models, where the

[23] Namely, it would be difficult to rationalise the short-run variability of the RER with reference to real shocks only, because they are not so frequent and, in any case, would tend to induce permanent deviations. On the other hand, it would also be difficult to attribute RER behaviour to the effect of nominal shocks, because their effect would be apparent for a short period of time, only (Rogoff, 1996).

[24] More recently, a promising path for research on the RER stylised facts has been opened by Moore and Roche (2006). They introduce a two-country monetary model, where preferences depend on an aggregate consumption externality (externally generated habit). This extension generates enough volatility and persistence to match the stylised facts of the RER.

[25] In order to overcome the low power problem, other strands of the literature adopted long span studies or panel unit root studies in linear settings (Abuaf and Jorion, 1990; Lothian and Taylor, 1996; Taylor, 2002). Although both methods provided supportive evidence of the PPP condition, it is still contentious whether favourable outcomes using these methods are enough to validate PPP (Frankel and Rose, 1996; Lothian, 1997; Taylor and Sarno, 1998; Sarno and Taylor, 2003)

RER is allowed to display different behaviour and, therefore, assume different speeds of adjustment at the different states. Several models were competing for the choice of the "correct" switching function, based on theoretical considerations about the nature of forces driving the RER behaviour (Leon and Najarian, 2005).

A big strand of non-linear unit root tests argues in favour of a discrete or smooth adjustment towards the PPP equilibrium, consistent with the limits to arbitrage theory. The latter relates to the existence of trade barriers and transaction costs (Dumas, 1992), which induce different dynamic adjustment of the RER towards its long run mean for different magnitudes of RER deviations from the PPP equilibrium. In case of discrete transition functions, fixed arbitrage costs create an implicit inaction band, within which the RER can float freely. The implication is that in this regime it is possible to observe a random walk in the RER. On crossing this threshold, however, arbitrage forces ensure that the RER process becomes mean reverting. Such behaviour is captured by a Threshold Autoregressive (TAR) model (Tong, 1990). Empirical application of a TAR model (Obstfeld and Taylor, 1997; Sarno, Taylor and Chowdhury, 2004, Leon and Najarian, 2005), provides support for the theory of discrete adjustments towards the PPP equilibrium and, thus, offers evidence in favour of the PPP. This empirical evidence, however, is overshadowed by scepticism over the width of the inaction band (see for example Kilian and Taylor, 2003) which has opened the way to alternative models.

However, advocates of smooth adjustment (Teräsvirta, 1994; Dumas, 1994; Bertola and Caballero, 1990) suggest a Smooth Transition Autoregressive (STAR) model (Teräsvirta, 1994) as an appropriate formulation[26]. This model assumes no explicit threshold, rather the speed of RER mean reversion to its long run equilibrium increases as the degree of misalignment from the PPP equilibrium increases. Further (simulation) analysis reveals

[26] Kilian and Taylor (2003) argue that a smooth adjustment of STAR type can also be due to the interaction of heterogenous agents in the foreign exchange market, namely economic fundamentalists, technical analysts and noise traders. As long as fundamentalists disagree about the level of the RER equilibrium, the traders will tend to act on information from the technical analysts. The latter follow trending techniques, which impart a unit root behaviour in the RER. However, as fundamentalists agree that the RER is far from its equilibrium, the tendency for the RER to revert back to its equilibrium is increasing.

that the mean reversion rate also varies with both the size of the RER shock and the initial conditions, that is the degree of RER disequilibrium when a given magnitude of shock hits the RER (Taylor, Peel and Sarno, 2001)[27]. Empirical applications of STAR model variants provide strong evidence of non-linear mean reverting behaviour for large deviations from the PPP equilibrium (Michael, Nobay and Peel, 1997; Taylor, Peel and Sarno, 2001).

Finally, a growing strand of literature is using Markov-Switching (MS) functions to model the behaviour of the RER. Such models allow for the distribution of the RER to be approximated as a mixture of normal distributions, and can, thus, permit changes in the speed of reversion, the mean and the variance of the RER process. Such models have been typically used for long-span data analysis, but RER applications with encouraging results are also found for the recent float (Leon and Najarian, 2005). The various regimes can depend on the deviation of the RER from its PPP equilibrium (Sarno and Valente, 2005), or the volatility of the RER shock (Engel and Kim, 1999) [28].

By allowing for different RER behaviour at the different states, the afore mentioned literature implicitly raised a further relevant question. This concerns the potentially different speeds of adjustment for positive or negative deviations of the RER from its PPP equilibrium, i.e. the possibility of asymmetric mean reversion towards the RER equilibrium. There is a considerable division of feelings in the literature over this issue, as theoretical and empirical arguments can be found in support for both sides. On the one hand, if goods arbitrage is driving the impetus back towards the long run PPP equilibrium, it would be difficult to explain why the speed of adjustment should be

[27] Notably, Taylor, Peel and Sarno (2001) provide further insight into the mean reverting process of the estimated non-linear (STAR) model, through a dynamic stochastic simulation. This allows the analysis of impulse response functions, where arbitrary magnitudes of shocks are imposed to drive the RER away from its equilibrium, in order to study the mean reverting path back to it and calculate half lives of shocks. Their findings suggest that if a shock of a given magnitude hits the RER and drives it further away from the equilibrium, the larger the shock, the faster the RER mean reversion. In that case half lives can fall just under one year (10 months).

[28] An extensive amount of literature has found that MS models are suitable for modelling the exchange rate behaviour (e.g. see Engel and Hamilton, 1990; LeBaron, 1992; Engel, 1994; Dueker and Neely, 2005 and Sarno and Valente, 2006).

different above or below the equilibrium (Taylor, Peel and Sarno, 2001). That is because, the limits to arbitrage theory, which motivates the specification of the TAR and STAR models, relies on the existence of symmetric transactions costs, and would, therefore, also require symmetric adjustment above or below the PPP equilibrium (Obstfeld and Taylor, 1997; Michael, Nobay and Peel, 1997; Taylor, Peel and Sarno, 2001, Sarno, Taylor and Chowdhury, 2004; Sarno and Valente, 2005).

On the other hand, a more recent strand of literature suggests that the limit to arbitrage theory cannot alone explain the dynamics of the RER. They bring forward the role of central bank intervention as an underlying force affecting the dynamic adjustment of the RER. In this context, asymmetries may arise as a result of intervention policies directed at the RER. Almekinders and Eijffinger (1996), provide evidence in favour of asymmetries in the intervention policies of the US and the German central banks in the post-Louvre period, by showing that the banks tried to counteract appreciations of their currency more strongly that depreciations. Taylor (2004) shows that net intervention from the same central banks could stabilise the RER, with the effect becoming bigger, the bigger the deviations of the RER from its equilibrium value. On the same note, Dutta and Leon (2002) argue that governments might want to defend an appreciation of the currency more or less rigorously than a depreciation, therefore inducing asymmetric dynamic adjustment behaviour. Finally, Leon and Najarian (2005) provide direct empirical support for the existence of asymmetries in the RER mean reverting behaviour across a wide range of countries[29].

In this chapter, we address the issues confounding previous PPP tests and also assess the symmetric properties of the RER mean reverting behaviour with the aid of the recently developed methodology of quantile unit root inference. Our unit root test adopts an agnostic approach towards the potential RER distribution and allows the RER

[29] Leon and Najarian (2005) adopt both a time-varying TAR model and a smooth transition (STR) model. In the first case, the magnitudes, frequencies and durations of the deviations of the RER from its forecast are allowed to differ for depreciations and appreciations. In the case of the STR model asymmetric adjustment is allowed for middle and outer regimes.

to assume different speeds of adjustment at different states, while naturally revealing asymmetries in the RER mean reversion process. As a result, the quantile unit root test provides an alternative approach for robust unit root inference. Our method effectively addresses the two PPP puzzles and further refines and enhances previous results in the PPP literature.

3.1.2 The quantile approach to the PPP puzzles

We investigate three major currencies (UK pound, Japanese yen and Euro versus the US dollar) using a recently developed, unit root test for non-normal processes based on the quantile autoregression (QAR) approach in both semi-parametric[30] (Koenker and Xiao, 2004a,b) and non-parametric (Koenker, Ng and Portnoy; 1994) settings. By using the more robust quantile unit root alternative we aim to refine previous results and shed further light into the PPP puzzle.

Quantile regression estimation (Koenker and Basset, 1978) allows one to estimate and conduct inference on a whole range of conditional quantile functions, that is models where quantiles of the conditional distribution of the response variable are expressed as functions of explanatory variables[31]. By making no prior distributional assumptions, the quantile regression examines quantiles of the conditional distribution, in order to uncover different stochastic dependencies in the different quantiles. It, therefore, provides a more complete and nuanced picture of how covariates influence the location and shape of the entire response variable distribution (Koenker and Xiao, 2004a).

More specifically, we consider QAR models, where the autoregressive (slope) parameters may vary with quantiles. In the case of the RER, different solutions in distinct

[30]The method is semi-parametric, in that it only assumes a linear relationship between the dependent and explanatory variables, without making any distributional assumptions.

[31]This is in sharp contrast with the traditional conditional mean estimation procedure, which assumes normality in estimating a single measure of the conditional mean function. In cases of Gaussian distributions, the latter estimation method would adequately describe the whole conditional distribution and would, in fact, enjoy a certain optimality. Moreover, the coefficients of the dependent variables would be independent of the specified quantiles.

quantiles may be interpreted as differences in the mean reverting behaviour of the RER at various quantiles of the conditional distribution of the RER, that is at different magnitudes of RER shocks. In that case, bigger (positive or negative) shocks correspond to more extreme (high or low) quantiles[32]. As a consequence, the quantile unit root test is modified to incorporate the effects of various sizes of RER shocks (Koenker and Xiao, 2004b), and is, therefore, more robust compared to standard unit root models. Furthermore, QAR unit root inference can reveal different patterns of mean reverting behaviour for positive or negative shocks to the RER and, thus, naturally expose asymmetries in both the distribution of RER shocks and the impact of these shocks in the dynamic adjustment process of the RER to its long run equilibrium[33].

Seen in a different way, the linear QAR model captures state dependencies in a way comparable to, but different from a non-linear MS, TAR or STAR model. The linear QAR model adopts a different characterisation of states, by allowing for multiple discrete regimes, which are chosen on the basis of the conditional distribution of the RER (i.e. RER shocks). This procedure can effectively expose transient and/or permanent states (i.e. quantiles) in the RER adjustment process, thereby presenting a more compete and nuanced picture of the RER dynamic behaviour. In this sense, the linear QAR model bodes well with the spirit of the aforementioned non-linear models.

Overall, QAR inference has significant advantages in analysing dynamics and persistence in time series with non-Gaussian distributions and can, thus, provide a more robust alternative to the standard unit root tests, while sacrificing little efficiency under normality (Koenker and Xiao, 2004a,b). In the context of PPP, the QAR approach provides an alternative, robust way of looking at the validity of the PPP, while addressing the question of whether different magnitudes of shocks may generate different

[32] A more refined analysis of the notion of "mean reversion at the different quantiles" is offered in section 3.2.1.

[33] Although limit to arbitrage models typically impose uniform or symmetric behaviour (Obstfeld and Taylor, 1997; Sarno, Taylor and Chowdhury, 2004; Taylor, Peel and Sarno, 2001), evidence on non-linear asymmetric dynamic adjustment, due to government policies (i.e. intervention) has been recently emerging in the literature (Dutta and Leon, 2002; Leon and Najarian, 2005).

(symmetric or not) persistency patterns on the RER. Our application is, to the best of the author's knowledge, the first contribution of quantile regression in this context.

3.1.3 Contribution, main results and structure of the chapter

The QAR analysis provides original insights in the RER behaviour because of its general, yet flexible formulation. In contrast to previous, parametric designs, the quantile framework adopts a more general approach. It remains agnostic about the underlying distribution of the RER, and, consequently, in the treatment of the causes and specification of the dynamic adjustment of the RER to its long run equilibrium. In other words, we may obtain evidence of dynamic adjustment, consistent with the previous parametric (non-linear) literature, but without specifying the nature of the parametric (non-linear) relationship. In this way, the quantile approach is nesting assumptions and results from previous parametric models, in an a-theoretical way, thus circumventing the need to discriminate across different parametric model formulations.

The generality of the quantile model is well exploited by a flexible estimating framework, where the researcher is allowed to choose the quantiles under investigation, and, therefore, determine the level of detailed analysis that needs be undertaken. In the context of the RER, the above qualities allow insights into the following: a) We are able to detect how different sizes of shocks affect the RER speed of adjustment, (irrespective or not of the RER disequilibrium point when the shock hits the RER). The shocks analysed are actual, observed shocks, whose sizes are determined endogenously by the model. This offers an original view into the role of shocks on the RER and enriches anecdotal evidence from previous literature (Taylor, Peel and Sarno, 2001; Engel and Kim, 1999) b) The quantile method is able to reveal asymmetries in both the distribution of RER shocks and their impact on the RER mean reverting behaviour in a simple, intuitive and yet effective way. In this way, we shed more light to the relevant debate, by providing evidence using an original and relatively more simple model. c)

As a result from the above, the quantile unit root test is a more robust alternative in cases of non-gaussian innovations, compared to standard unit root tests. Overall, the quantile analysis sheds light into the two PPP puzzles by further refining and enhancing results previously obtained by, amongst others, Taylor, Peel and Sarno (2001), Leon and Najarian (2005).

More specifically, our results suggest that the RER is not a standard linear stationary or a constant unit root process. Namely, we find that: a) the dynamic behaviour of the RER is affected by the magnitude of RER shocks, with large RER shocks undermining the unit root behaviour of the RER and inducing potentially strong mean reverting tendencies. Half lives in that case can fall well below one year. b) When large shocks to the RER originate at large RER disequilibrium levels (i.e. far away from its PPP equilibrium), the effect can be even stronger. c) On the contrary, small shocks to the RER considerably weaken mean reversion tendencies, irrespective of the disequilibrium point of the RER at the time of the shock. d) There are marked asymmetries in the behaviour of the RER, i.e. extreme positive shocks can generate different reversion patterns than extreme negative shocks. Their extent also depends on the original condition of the RER with respect to its long run equilibrium.

The chapter proceeds as follows. Section 3.2 introduces the quantile regression techniques employed in this chapter. Section 3.3 describes the data and some preliminary data analysis. Section 3.4 presents the empirical results from the semi-parametric and non-parametric quantile approach, and Section 3.5 concludes.

3.2 Methodology

In this section we present the QAR framework in both its semi-parametric and non-parametric settings. We begin with the simple linear QAR(1) model and explain the estimation and inference procedure (i.e. quantile unit root tests within each quantile), as presented in Koenker and Xiao (2004b). We then proceed to a basic exposition of the non-parametric quantile estimation technique (Koenker, Ng and Portnoy, 1994).

The semi-parametric and non-parametric settings correspond to a general and a more refined analysis of shocks respectively. In the general (semi-parametric) analysis we consider different magnitudes of RER shocks in a linear QAR context in order to investigate their impact on the mean reversion of the RER. This methodology allows different speeds of adjustment for different magnitudes of RER shocks. However, this analysis does not consider the origin of the shock, i.e. the deviation of the RER from its RER long-run equilibrium when the shock occurs[34]. The limit to arbitrage theory offers plausible support for such considerations. We, therefore, further refine our results with a non-parametric quantile model. In that context, we observe patterns of RER behaviour, which are identifiable primarily by the magnitude of shocks (size), but also by the level of RER disequilibrium when the shock occurred (origin). We can, therefore, gauge results about different speeds of adjustment when shocks of given magnitude hits the RER on, below or above its equilibrium.

3.2.1 Semi-parametric QAR model

Our semi-parametric analysis is founded on the recent extension of the theory of quantile regression to autoregressive models, which resulted in the linear QAR model (Koenker and Xiao, 2004b). We use a linear QAR estimation framework on the deviation of the real exchange rate from its equilibrium value and perform different quantile unit root tests in order to gain a more refined view of the RER dynamic behaviour.

[34]Note that there is an important difference between RER shocks and RER deviations from equilibrium. A shock hits the RER at a time t and has an observable impact on the RER at time $t+j$. A shock is equal to a RER deviation if they both occur at the same time interval and if the shock originates at equilibrium. However, shocks conditional on the past history of the RER can occur at any point of the RER distribution with respect to the equilibrium (i.e. can occur when the RER is below or above its long run equilibrium). Because of that, RER deviations can be the additive result of cumulative shocks to the RER and the two expressions are no longer tautologous. Overall, the effects of shocks on the RER can be variable, depending on the magnitude of the shock and the RER disequilibrium position at the impact.

Estimation of the QAR model Let us consider a simple first order autoregressive, AR(1), model of the type

$$y_t = \alpha y_{t-1} + \varepsilon_t, \tag{3.1}$$

where $y_t = q_t - \mu$, with q_t denoting the logarithm of the RER and μ being the long run equilibrium level of q_t, i.e. the unconditional mean of q_t. Following the standard literature, the RER is defined as $q_t \equiv s_t - p_t + p_t^*$, where s_t is the logarithm of the nominal exchange rate (domestic price of foreign currency) and p_t and p_t^* denote the logarithms of the domestic and foreign price levels respectively. Hence, y_t represents the deviations of the real exchange rate from its equilibrium value. Finally, ε_t is an error term. In this traditional conditional mean function, standard unit root theory suggests the existence of a unit root in the RER, if the autoregressive coefficient, α, equals unity. In that case, deviations from the long run RER equilibrium are permanent. However, if the autoregressive coefficient is smaller than unity, the real exchange rate is a stationary process, suggesting that any deviations from equilibrium are transitory.

Following the methodology set out by Koenker and Xiao (2004b), the equivalent τ^{th} quantile representation takes the form:

$$Q_{y_t}(\tau \mid y_{t-1}) = Q_{\varepsilon_t}(\tau) + \alpha(\tau) y_{t-1}, \tag{3.2}$$

where $Q_{y_t}(\tau \mid y_{t-1})$ is the τ^{th} conditional quantile of y_t, conditional on y_{t-1}, and $Q_{\varepsilon_t}(\tau)$ is the τ^{th} conditional quantile of ε_t. In other words, the τ^{th} conditional quantile function of the dependent variable y_t is expressed as a linear function of its own lagged value. $\alpha(\tau)$ is the autoregressive coefficient, which measures the persistence of the real exchange rate deviations within each quantile and is dependent on the τ^{th} quantile under investigation.

Estimation of the linear QAR model involves solving a minimisation problem of weighted residuals, where all the observations are considered, but are being weighted in such a way, so that the residuals fall into the selected quantile:

$$\min_{\alpha\in\mathbb{R}^2} \sum_{t:y_t\geq x_t^\intercal \alpha} \rho_t(y_t - x_t^\intercal \alpha(\tau)), \tag{3.3}$$

where $\rho_t(\varepsilon) = \varepsilon(\tau - I(\varepsilon < 0))$ is a check function with I denoting an indicator taking the value of 1 if the expression in parentheses is true and 0 otherwise, $x_t = (1, y_{t-1})$ and $\alpha(\tau) = (Q_\varepsilon(\tau), \alpha(\tau))$. Thus, equation (3.3) is equivalent to:

$$\min_{\alpha\in\mathbb{R}^2} \sum_{t:y_t\geq x_t^\intercal \alpha(\tau)} \tau(y_t - x_t^\intercal \alpha(\tau)) + \sum_{t:y_t< x_t^\intercal \alpha(\tau)} (\tau - 1)(y_t - x_t^\intercal \alpha(\tau)). \tag{3.4}$$

In our case the QAR model was estimated in the "quantreg" package included in R, using a modified simplex algorithm of Barrodale and Roberts (Koenker and d'Orey, 1987, 1994). This package offers the possibility to estimate a whole range of conditional quantile functions and computes bootstrapped standard errors for the parameters. In our case the number or replications employed were 2000.

Quantile unit root tests A general analysis of the unit root behaviour based on the quantile approach involves examining the unit root property over a range of quantiles. The relevant statistic for testing the null of a constant unit root process over a range of quantiles is a Kolmogorov-Smirnov (KS) test based on the regression quantile process over a range of $\tau \in T$. Koenker and Xiao (2004a,b) suggest

$$QKS = \sup |t(\tau)|, \tag{3.5}$$

where $t(\tau)$ is the t-statistic of the autoregressive coefficient at the τ^{th} quantile. In practice, we may calculate $t(\tau)$ at $\tau \in T$ and construct the QKS statistic by taking the maximum statistic value over $\tau \in T$. The limiting distribution can be approximated by resampling methods, as explained below.

A more detailed examination of the unit root properties of the series is by examining the unit root property in each quantile separately. This allows for a closer look at the dynamics of the series and also permits the detection of possible asymmetries in the process. The relevant unit root test involves a simple t-statistic test, $t(\tau)$ for the null of a unit root. In other words, we are testing that the autoregressive coefficient in the specific quantile, $\alpha(\tau)$, will be equal to unity. Given that $\alpha(\tau)$ depends on τ, it is possible to have different mean reverting behaviour in the different quantiles. This implies that it is possible to observe sequences of innovations that reinforce the unit root behaviour of the series, followed by occasional realisations that induce mean reversion and thus undermine the persistency of the whole process.

For both types of tests we base our inference on a resampling (bootstrap) exercise, as described by Koenker and Xiao (2004b)[35], which was coded in R. The main idea of this exercise is to generate a distribution for the relevant statistic values and observe where our actual statistic values lies with respect to the bootstrapped distribution. For this purpose, we construct dependent variables (y_t) under the null of a unit root in the RER data generating process, by resampling from the original data. We then estimate the same quantile regression specification under the null and get the relevant t-statistic values. We repeat this procedure 2000 times. We, thus, create the distribution of the $t(\tau)$ test and generate the distribution of the QKS. We then compare the statistic value of the original (true) regression with the distribution under the null (of a unit root). The percentage amount of times that the statistic value will be above the bootstrapped statistic value gives us the probability of rejecting the null hypothesis of a unit root, within each quantile. In this study, we investigate a range of quantiles for $\tau = (0.01, 0.05, 0.1, 0.25, 0.5, 0.75, 0.9, 0.95, 0.99)$[36].

Koenker and Basset (2004b) by means of a Monte-Carlo analysis, compare the power

[35] Methods of asymptotic inference are also available for the $t(\tau)$ test. The asymptotic distribution is not the conventional Dickey–Fuller distribution, but rather a linear combination of the Dickey–Fuller distribution and the standard normal.

[36] For an explicit technical description of the procedure see Koenker and Xiao (2004b).

of the OLS and QAR models for the case of Gaussian and Student-t innovations. Their results show that the quantile-based tests have superior power than the simple Augmented Dickey-Fuller (ADF) and Phillips-Perron tests in cases of Student-t innovations. In turn, the $t(\tau)$ test has more power than the QKS test, albeit marginally.

Interpretation of the quantiles and quantile mean reversion In order to interpret our results from the QAR model, it is important to consider first the meaning of each quantile, i.e. what exactly the quantiles capture, and second the meaning of quantile mean reversion. As regards the first issue, looking at the QAR specification and the estimation procedure (equations 3.2 – 3.4), it becomes obvious that the quantile approach estimates quantiles of the conditional distribution of the RER, conditional on its own past values, i.e. it estimates quantiles of the error term. Therefore, in the simple case of a QAR(1) model the quantiles capture the magnitude of shocks from period $t-1$ to period t[37]. That is, one-off shocks of similar magnitude, which are classified as falling into the same quantile are, in effect, the shocks that determine the fit of this quantile. The magnitude of these shocks is summarised by the constant term, $Q_\epsilon(\tau)$. Therefore, the more extreme the quantile the more extreme the shocks that hit the RER in the same quantile.

The quantile methodology has the potential to reveal different localised mean reverting patterns, by explicitly testing for a unit root at the different quantiles (i.e. locally). More specifically, RER mean reversion at a specific quantile suggests that shocks of similar magnitude, that fall into this quantile, tend to undermine the persistency of the series and induce mean reversion tendencies on the RER. On the contrary, unit root behaviour within a quantile suggests the existence of innovations of a certain magnitude, which reinforce the persistency of the RER. It is, therefore, possible for a series to exhibit localised unit root behaviour (i.e. unit root in certain quantiles), followed by

[37] In a case of a higher order QAR model, of the type specified by Koenker and Xiao (2004a,b), they would capture the cumulative effect of the $t-n$ periods to period t, where n is the number of lags allowed for.

mean reverting occasions (i.e. mean reversion in other quantiles) capable of inducing stationarity in the overall process (i.e. globally).

3.2.2 Non-parametric QAR estimation

In the next step of our analysis we move to a non-parametric QAR framework. Namely, we investigate if the impact of different magnitudes of RER shocks is further affected by initial conditions (i.e. the level of RER disequilibrium when the shock hits the RER). According to the limits to arbitrage argument, should large deviations from the PPP equilibrium affect mean reversion, then the linear fit should not be a good approximation of the quantile process and instead we should observe kinks (i.e. different slopes) in each of the different quantile fits. We, therefore, employ a non-parametric model in an effort to allow for a more flexible functional form within each quantile compared to a semi-parametric one. Our aim is to expose distinct linear sub-segments, i.e. linear sub-segments with different gradients, within each quantile.

The preferred non-parametric estimation technique, is the method of quantile smoothing splines with total variation roughness penalty (Koenker, Ng and Portnoy, 1994). If y, x and τ are defined as above, the idea underlying this method is to derive the quantile smoothing spline estimator of $g_\tau(x)$, as the solution to a trade-off problem between "fidelity" and "roughness", i.e. between a fit the bears a reasonable degree of fidelity to the observed points and a fit with a plausible degree of smoothness:

$$\min_{g} \text{“}fidelity\text{”} - \lambda \text{“}roughness\text{”} \tag{3.6}$$

where,

$$\text{“}fidelity\text{”} = \sum_{y_i - g(x_i) \geqslant 0}^{n} \tau(y_i - g(x_i)) + \sum_{y_i - g(x_i) < 0}^{n} (\tau - 1)(y_i - g(x_i)) \tag{3.7}$$

and

$$\text{“}roughness\text{”} = V(g'). \tag{3.8}$$

Therefore, the quantile smoothing spline estimator is the solution to

$$\min \sum_{i=1}^{n} \rho_{\tau}(y_i - g(x_i)) - \lambda V(g') \tag{3.9}$$

where g is a smooth function with a uniformly continuous first derivative g' and bounded second derivative g''. In our approach, lambda (λ) penalises the total variation of function g', which we denote as $V(g')$, with $V(g') = \int_a^b |g''(x)|\, dx$. λ is a regularisation parameter, or the roughness penalty, that balances the trade off between fidelity and roughness and therefore determines the smoothness of the fitted function. As λ increases the penalty prevails until, for very high values of λ, the roughness penalty is maximised and we get a perfectly smooth line, matching the semi-parametric linear fit. The solutions are piecewise linear functions with knots at x_i (Koenker, 2005).

The estimation techniques for this type of non-parametric fit depend on the dimensionality of the vector of conditioning variables x. Our QAR(1) case corresponds to a univariate case of non-parametric smoothing and for this purpose the quantile model was estimated using the COBS (Constrained B-Splines Smoothing) algorithm of He and Ng (1999)[38].

The quantile smoothing spline methodology is appealing in our case, both technically and intuitively, since it provides a direct comparison with the semi-parametric linear fit, while also allowing a role for the limits to arbitrage theory. Namely, in each quantile we are testing the robustness of the linear fit, thereby investigating the validity of the limits to arbitrage argument. In the quantiles where deviations from the RER do not affect the mean reversion properties of the RER, the graphical results should deliver the same linear quantile fit found using the semi-parametric methodology. In the opposite case,

[38]The COBS package in R permits the implementation of this algorithm and also enables the calculation of confidence intervals, based on the asymptotic results of He and Shi (1998).

however, where large RER deviations from its equilibrium value induce mean reversion, the linear quantile fit should change to a piecewise linear quantile fit. If the limits to arbitrage theory is supported by our data, we would expect to find sub-segments with less than unity slope at the left and right hand side of a particular quantile, while the middle part could preserve a unity slope. Such a result would suggest that, when a given magnitude of shock is originated at a high disequilibrium level, the mean reversion of the RER is stronger.

3.3 Data

The data sources used to construct our RER data set are the International Monetary Fund (IMF)'s, International Financial Statistics (IFS) and the Organisation for Economic Co-operation and Development (OECD)'s, Main Economic Indicators (MEI). The countries analysed include the euro area, the UK and Japan with the US as the reference country, for a period from January 1973 to December 2004. For each country, we obtained the relevant nominal bilateral (end-of-period) exchange rates vis-à-vis the US dollar. These were the euro (EU), the UK pound (GBP) and the Japanese yen (JY) denominated in US dollar (USD) terms. In order to prolong the EU nominal exchange rate series, euro-dollar values before the introduction of the euro were proxied with Deutsche mark-dollar data. CPI (total index) monthly data for the five countries were collected from MEI. The final times series - monthly RER (deviations from the long run mean) in logarithmic terms (y_t) - were constructed following the RER formula mentioned in Section 3.2.1.

As a preliminary exercise we compared the sample moments of the RER deviations of the three exchange rates in question, and performed normality tests (Table 3.1). For the individual series in levels, the summary statistics confirm evidence of leptokurtosis and non-normality. The formal Jarque-Bera test rejects normality in every case, adding support for using quantile regression.

3.4 Empirical results

In this section we report estimation results from the semi-parametric linear QAR model and the non-parametric quantile smoothing method. Results are complemented with calculations of the relevant half lives. The semi-parametric method provides some evidence of mean reversion across a range of quantiles. A more detailed and instructive view is taken by focusing on the specific quantiles, where the mean reversion becomes much stronger in the extreme quantiles (i.e. for extreme RER shocks). The non-parametric test further reveals that the behaviour in each quantile is exacerbated when extreme shocks combine with extreme deviations from the RER long run equilibrium.

3.4.1 Estimation, unit root tests and half lives

As a first step, we had to choose the order of the AR process. Towards this, we followed previous practices from Granger and Teräsvirta (1993), Teräsvirta (1994) and Taylor, Peel and Sarno (2001) and focused on the partial autocorrelation function. In our case (results not reported, but available upon request) this analysis clearly reveals that only the first partial autocorrelation coefficient is significantly different from zero at the five percent level. Overall, in all cases a simple AR(1) model sufficiently captures the dynamics involved. We enhance this result with a test for residual correlation (F_{RS} test in Table 3.2, Panel A), where we find that we can reject the hypothesis of serial correlation at the five percent level for the AR(1) specification. We then estimate a conventional conditional mean specification, i.e. a simple AR(1) model using OLS, and a QAR(1) model for $\tau = (0.01, 0.05, 0.1, 0.25, 0.5, 0.75, 0.9, 0.95, 0.99)$. For both specifications we performed (quantile) unit root tests (Table 3.2, Panels A and B). Our analysis is completed with the estimation of half lives[39] for both the AR(1) and the QAR(1) models (Table 3.3, Panels A and B).

[39]Half lives for a simple AR(1) model are computed based on the formula $\log(0.5)/\log(a)$, where α is the autoregressive coefficient under consideration.

Conditional mean (OLS) specification results The naive conditional mean estimate of the autoregressive coefficient in the AR(1) model confirms the stylised facts of a unit root in the RER, with all estimated coefficients very close to unity and the relevant unit root tests in levels suggesting that the coefficient values are not statistically different from unity at the five percent level of significance. Overall, evidence from the two unit root tests employed, the Phillips-Perron (1988) and the Ng and Perron (2001) tests, on the levels and first differences indicate that, while changes in the real exchange rate are stationary, the level of the real exchange rate contains a unit root. These findings replicate well established results in the literature.

QAR specification results We now reconsider these series using quantile unit root tests. In particular, we first apply the QKS test based on the QAR model for a range of quantiles $T = (0.01, 0.99)$. This test gives us a general idea of the unit root behaviour of the series in question. Results are reported in Table 3.2, Panel B. Contrary to the conventional unit root tests presented above, QAR unit root tests provide some evidence in favour of mean reversion for the GBP and the EU, at the 10% significance level. For the JY, we cannot reject the null of a unit root. This is not overall unexpected for the JY, which has, in fact been notorious for such type of behaviour. This could be the result of the Japanese catching up after the WWII, creating productivity differentials which determine the long run equilibrium of the RER (Harrold-Balassa-Samuelson effects).

By and large we find the comparison between the two sets of results encouraging. For a more refined investigation, we turn our attention towards the behaviour of specific quantiles (Table 3.2, Panel B). The first striking observation is a varied behaviour across the different quantiles, both for the intercept (τ_0) and the autoregressive coefficient ($\alpha(\tau)$). As noted above, the intercept captures the magnitude of the typical, observed RER shock in each quantile (negative signs suggesting negative shocks, loosely interpreted as appreciations and positive signs suggesting positive shocks, loosely interpreted as depreciations). The τ_0 coefficients present a monotonically ascending, symmetric behaviour, i.e. the absolute magnitudes of positive and negative RER shocks are quite similar for a given set of complementary (symmetric) quantiles (e.g. the 1% and the 99% or the 25% and 75% quantiles), therefore the magnitudes of shocks hitting the RER appear to be symmetric. We also observe that the magnitudes of shocks are similar across currencies, with the biggest shocks in absolute value deviating from the long run equilibrium by approximately 0.035 log units.

However, the most interesting results are the values of the autoregressive (slope) coefficients $\alpha(\tau)$ and the relevant unit root tests in the QAR(1) model, which determine the mean reverting behaviour of the RER in each quantile. A careful look reveals a distinct pattern and gives clear support for mean reversion in certain quantiles and unit root behaviour in others. In particular, in the middle quantiles we observe coefficient values very close to unity (above and below), and in fact not different from unity in statistical terms, suggesting a unit root in the RER[40]. However, in the extreme quantiles coefficients appear to be lower with p-values rejecting the null of a unit root in conventional significance levels, suggesting that the persistence in the RER drops. A graphical representation of the above results is produced in Figure 3.1, where the values of the autoregressive coefficient for the different quantiles are displayed. It is possible to see an inverse U-shaped pattern, suggesting that the coefficient is smaller in the extreme quantiles than in the mean quantiles. This heterogeneity in the slope coefficients suggests a dynamic adjustment towards the long-run PPP equilibrium. In fact, our main conclusion is that, in the presence of small and medium shocks, the RER does not adjust towards its PPP equilibrium value, but extreme shocks seem to have the potential to induce mean reversion. These results are in line with evidence from Taylor, Peel and Sarno (2001) and also relate to relevant evidence from Engel and Kim (1999).

Our results from the $t(\tau)$ test identify that the series under consideration are not constant unit root processes. Nevertheless, the inconsistency with the QKS test, for

[40]It is also worth noting that the estimated autoregressive coefficient (α) in the conditional mean model assumes values very close to the conditional median quantile estimates.

the case of the JY raises questions about the global properties of this series. This inconsistency could be due to the lower power of the QKS test. However, our series do not follow a Student-t distribution, therefore evidence on the comparison of the two unit root tests in terms of power (Koenker and Xiao, 2004b) is weak. Overall, our results suggest that there are cases (i.e. quantiles) where the RER is mean reverting, and these tend to be cases where big shocks hit the RER (i.e. the most extreme quantiles). We can say with some certainty (90%) that this is enough for the whole process to revert back to its long run mean (apart from the case of the JY).

Turning our attention to the estimated half lives (Table 3.3, Panels A and B), in the simple AR(1) model half lives are equal to infinity, because a unit root behaviour dominates the results. However, in the QAR(1) model, for the mean reverting quantiles we get different results. Namely, in the very extreme quantiles (99%) we get surprisingly low half lives, ranging from 5 to 8 months. Half lives increase, but still remain quite low in the 95% quantile, ranging from 10 to 14 months and only in the 90% quantile we can see half lives of more than one year. These findings are well below the four year average suggested by Rogoff (1996) and also below the findings of the non-linear literature.

We, therefore, see that the simple linear quantile model, in its ability to conduct analysis on the different magnitudes of RER shocks, can give signs of mean reversion at the different quantiles, consistent with the PPP.

Asymmetric dynamics It is interesting, however, to note that only in the case of the GBP this effect appears symmetric, i.e. the RER is a less persistent process in both extreme positive and negative shocks (although more so for extreme positive shocks). In the case of the EU and the JY the RER appears to be mean reverting only in cases of extreme positive shocks. In statistical terms, this asymmetry suggests a shortage of extreme values in the low or high quantiles with the potential to induce mean reversion. This might be the case either because extreme shocks do not occur or because the shocks of different signs weight differently. Given the symmetric magnitudes of shocks, as reported from the constant term values (τ_0), it is more plausible to assume the latter explanation. A potential reason for such asymmetries might lie on monetary policy choices and official intervention that impact asymmetrically on exchange rates (Dutta and Leon,2002; Leon and Najarian, 2005). Overall, the semi-parametric quantile analysis suggests that, although the shocks that hit the RER are of symmetric magnitudes, they impact asymmetrically on the mean reversion of the RER.

3.4.2 A graphical representation of the RER behaviour

A closer look at Figure 3.2, should provide a clearer intuition on the focus and results of the QAR(1) model. In the graph we plot the realisations of the RER on the lagged value of the RER for the GBP. Given that our data are monthly, the graph plots the realisations (dots) of the UK pound RER this month against its value in the previous month. The straight diagonal line is the 45 degree, $x = y$ axis, suggesting that the RER has not changed since last month, that is, shocks on the RER are practically zero. This further implies an autoregressive coefficient of unity and therefore a unit root process. All the dots above the diagonal line suggest negative shocks (depreciations) to the RER, because a deviation at time t is followed by a bigger deviation at time $t+1$. Alternatively, all realisations below the diagonal suggest positive shocks (appreciation), because a deviation at time t is followed by a smaller one at time $t+1$. The further away we move from the diagonal line, the bigger the shock becomes, be it positive or negative.

A first look at the realisations can give a deceptive unit root impression, since most realisations lie across the diagonal line. A closer look, however, reveals different patterns, namely that the centre of the graph is more dense, i.e. most realisations lie close and around the long run equilibrium, whereas the tails of the unconditional distribution (top right hand and bottom left hand corner) are not only more sparsely populated but also have relatively bigger deviations from equilibrium, i.e. there appear to be either large positive or negative shocks.

The dotted line is the mean (OLS) and the dashed line is the median (50% quantile) fit. It is obvious that the slopes of both lines are very similar to each other and are, in fact, difficult to discern from the diagonal (long-slash) line, suggesting that the OLS and median quantile outcome will favour a unit root behaviour. However, the image changes when we look at the outer slashed lines, which are the fits on the 1% quantile (lowest line) and the 99% quantile (highest line), representing extreme negative and positive shocks to the RER. In our case, the slopes of the extreme quantile fits are definitely smaller than the slope of the diagonal, suggesting that extreme shocks tend to induce mean reversion in the RER.

Finally, it is important to note that, for a given quantile, the slope is determined by RER realisations that are close to (points in the middle part of the quantile fitted line) or far away from (points at the two ends of the quantile fitted line) the PPP equilibrium. That is, each quantile fit depends on realisations (shocks) that hit the RER at various RER points with respect to its PPP equilibrium. A linear fit suggests that it is only the magnitude of the shocks and not the original conditions of the RER at the time of the shock that affect the fit. However, the limit to arbitrage theory suggests that original conditions can impact on the mean reversion of the RER. In order not to ignore potentially richer dynamics, that might result when a shock occurs far away from the PPP equilibrium, we accommodate such considerations in the non-parametric part of our analysis.

3.4.3 Non-parametric results

In this section we present the results of the piecewise linear fit, obtained by non-parametric quantile smoothing, using total variation regularisation, following the methodology set out by Koenker, Ng and Portnoy (1994)[41]. A graphical representation of the

[41] The choice of lambda, and thus the number of distinct linear segments, was based on the minimisation of a modified Schwartz information criterion (SIC).

results is presented in Figure 3.3, Panels A to C. For each currency we impose the 1%, 50% and 99% quantiles of the piecewise linear fit on a line with unity slope and a constant equal to the respective quantile τ_0 coefficient, so that any discrepancy between the unit root case and the non-parametric fit is easier to detect. Figure 3.4 (Panels A to C) graphically presents the various slopes in the individual quantiles under consideration, corresponding to the piecewise linear plots in Figure 3.3. Finally, the analysis is complemented with Table 3.3, Panel C, where we show the slope coefficients and the relevant half lives for each subsegment of the piecewise linear fit.

Looking at the extreme quantiles we observe distinct departures from the linear QAR model. The multiplicity of linear sub-segments within the same quantile stresses the difference between the semi-parametric and the non-parametric method and, moreover, offers support to the limits to arbitrage theory. A careful look will reveal that the left and right end of the 99% and 1% quantiles are, in the majority of cases, associated with strong mean reverting RER behaviour for all currencies involved. Figure 3.4 and Panel C of Table 3.3 gives ample support to that observation, with half lives recording very fast mean reversion, as low as 1.3 months (GBP in 1% quantile). This outcome is much stronger compared to the previous results of the literature, and even stronger than our results in the semi-parametric model. For the middle part of the extreme quantile fits, however, we get evidence of an autoregressive coefficient close to unity in most cases. In line with the limits to arbitrage argument, evidence from the extreme quantiles suggests that large shocks, which originate at large disequilibrium levels, tend to induce strong RER mean reversion. Mean reversion tendencies in the presence of large shocks are much weaker around the RER long run equilibrium.

A quite novel insight comes from looking at the behaviour of the median quantiles, which differs significantly from the one mentioned above. In the median (50% quantile) the fit appears to be the same as in the linear case. Note that in the median quantiles the shock to the RER is minimal. This leads us to conclude that, in the absence of shocks, the dynamic behaviour of the RER is not affected, irrespective of the RER deviation from the equilibrium.

As regards asymmetric dynamic adjustment patterns, compared to the linear, semi-parametric QAR model, we find evidence of mean reversion in both extreme quantiles for all currencies, although by no means exactly symmetric. However, a more careful examination reveals a pattern. Asymmetries in the adjustment dynamics of the RER are more pronounced when large shocks hit the RER at points far away from its equilibrium. Asymmetries become less pronounced, or even disappear when large shocks hit the RER near its equilibrium value. Finally, in the absence of shocks, for any disequilibrium level, we cannot establish asymmetric dynamic adjustment patterns.

By and large, our results in this section offer support to the limits to arbitrage theory, put forth by non-linear TAR and STAR methodologies. In the mean time, by taking into account both the effect of different magnitudes of RER shocks and the original disequilibrium condition of the RER we manage to find half lives significantly smaller compared to the previous literature. Overall, our results suggest the following about the driving forces behind the RER mean reverting behaviour: a) When a big shock hits the RER at a point already far from its equilibrium level, this shock tends to induce mean reversion. b) Big shocks that originate at points near the PPP equilibrium have much reduced mean reversion abilities. c) Small shocks either around or away from the RER equilibrium do not appear to induce mean reversion. d) We find asymmetries in the adjustment dynamics of the RER when large positive or negative shocks of the same absolute magnitude hit the RER at large disequilibrium points. Asymmetries become less pronounced when a big shock hits the RER near its equilibrium value or in the absence of shocks, despite the disequilibrium level of the RER.

3.5 Conclusions

This chapter elaborates on the long standing PPP puzzles. Earlier literature seeked answers by employing unit root tests with different levels of sophistication. Amongst those, the ones which accommodated the non-Gaussian behaviour of the RER, seemed to have better power in detecting reversion towards the PPP equilibrium. In this chapter, we present QAR semi-parametric and non-parametric methods as an alternative approach for robust inference in non-gaussian series. The quantile approach adopts an agnostic and yet flexible framework for the analysis of the RER behaviour, thus sidestepping the need to specify theory-consistent driving forces of the RER dynamic adjustment process. More precisely, the quantile framework makes no assumptions about the underlying distribution of the RER, while allowing for different (symmetric or asymmetric) persistence patterns at the different quantiles. In this sense, it is possible to observe sequences of unit-root behaviour, while occasional mean reverting tendencies can undermine the persistence of the whole process. By taking into account the different adjustment processes at the different quantiles, the quantile approach offers a more robust unit root test than standard alternatives.

More importantly, the QAR analysis and inference sheds light into both PPP puzzles. As concerns the first one, our methodology offers some support for the PPP, by providing evidence in favour of a mean reversion in the RER from two different quantile unit root tests. Our approach also addresses the second PPP puzzle by undertaking a detailed analysis of the impact of different magnitudes of actual shocks on the RER. We rationalise the high persistence of the RER behaviour, by suggesting that different magnitudes of shocks can induce different speeds of adjustment to the RER, while maintaining consistency to the limit to arbitrage theory.

More specifically, our evidence from two different quantile unit root tests in semi- and non-parametric settings suggests that the RER is not a constant unit root process across quantiles. We find that the bigger the shock to the RER (i.e. the higher the quantile) the faster the mean reversion back towards its long run equilibrium, with half lives comfortably less than a year, in the case of extreme shocks. Our results are further enhanced when large shocks hit the RER at points already far from its equilibrium. In such cases half lives can fall significantly less than a year. However, the mean reversion ability of large shocks is diminished in cases when the RER is around its equilibrium value. Finally, in the absence of shocks, mean reversion cannot be established irrespective of the RER disequilibrium level. In addition, our method captures asymmetric dynamic adjustment of the RER, i.e. positive shocks have different impact than negative shocks. Our results offer novel insights on the RER mean reverting behaviour and further refine and enhance previous evidence in the PPP literature.

Table 3.1 RER descriptive statistics

	GBP	JPY	EUR
Mean	1.04E-10	-1.30E-11	1.08E-02
St. Dev.	0.057	0.093	0.143
Skewness	0.325	0.061	-0.538
Kyrtosis	3.251	2.108	2.055
Jarque Bera	7.829 [0.019]**	13.178 [0.001]*	33.247 [0.000]*

Notes. The table presents the results from the descriptive statistics (the first four moments) and the Jarque-Bera normality test for the logged values of the RERs analysed. Values in brackets are asymptotic p-values. One and two asterisks denote significance in the 1% and 5% level respectively.

Table 3.2 Autoregression estimation and unit root tests

Panel A) Conditional mean (OLS) specification (lags=1)

OLS	GBP	JPY	EUR
α	0.974 (0.012)	0.981 (0.018)	0.988 (0.005)
F_{RS}	[0.089]*	[0.087]*	[0.088]*
	Unit root tests in levels		
PP	−2.276 [0.180]	−2.394 [0.144]	−2.322 [0.165]
MZ_a	-2.954	-0.474	-0.614
	Unit root tests in first differences		
PP	−18.145 [0.000]***	−18.091 [0.000]***	−18.245 [0.000]***
MZ_a	-78.483	-20.594	-23.139

(continued...)

(...Table 3.2 continued)

Panel B) Quantile autoregressive linear specification (lags=1)

Quantile	GBP		JPY		EUR	
QKS unit root test over a range of quantiles (1% to 99%)						
1%-99%	[0.073)]		[0.492]		0.093	
Quantile estimation and unit root tests within each quantile						
	τ_0	$a(\tau)$	τ_0	$a(\tau)$	τ_0	$a(\tau)$
1%	−0.034 [0.000]***	0.875 [0.023]**	−0.039 [0.000]***	0.960 [0.138]	−0.035 [0.000]***	0.992 [0.541]
5%	−0.022 [0.000]***	0.958 [0.455]	−0.028 [0.000]***	0.985 [0.600]	−0.021 [0.000]***	0.988 [0.261]
10%	−0.017 [0.000]***	0.964 [0.292]	−0.020 [0.000]***	0.986 [0.671]	−0.014 [0.000]***	0.988 [0.359]
25%	−0.009 [0.000]***	0.987 [0.715]	−0.008 [0.000]***	0.986 [0.455]	−0.007 [0.000]***	0.994 [0.680]
50%	−0.000 [0.472]	0.988 [0.690]	0.000 [0.349]	0.992 [0.292]	0.001 [0.070]*	0.993 [0.427]
75%	0.008 [0.000]***	1.009 [0.957]	0.008 [0.000]***	0.984 [0.179]	0.010 [0.000]***	0.992 [0.461]
90%	0.016 [0.000]***	0.984 [0.613]	0.016 [0.000]***	0.977 [0.079]*	0.020 [0.000]***	0.973 [0.283]
95%	0.021 [0.000]***	0.953 [0.075]*	0.021 [0.000]***	0.953 [0.065]*	0.027 [0.000]***	0.935 [0.060]*
99%	0.032 [0.000]***	0.898 [0.010]***	0.032 [0.000]***	0.912 [0.055]*	0.034 [0.000]***	0.918 [0.015]**

Notes: *Panel A)* The table shows the estimated values of the autoregressive (α) coefficient of a simple AR(1) model with the correspondent standard errors in parenthesis, the p-values of the residual correlation F-test (F_{RS}), and two unit root tests, the Philips-Perron (PP) (test statistic and p-values in brackets) and the Ng and Perron (MZ_a) test statistic. Unit root tests are reported for the level and the first difference of each series in question. The critical values for the MZ_a test are -13.800, -8.100 and -5.700 for the 1%, 5% and 10% level respectively. *Panel B)* The table shows the bootstrapped p-values (2000 replications), calculated using the pair-wise bootstrap method for the Kolmogorov-Smirnov (QKS) test, for the null of a unit root over a range of quantiles $\tau \in T$, where $T = (0.1, 0.99)$. It also shows the estimated values of the constant term (τ_0) and autoregressive ($a(\tau)$) coefficient of a QAR(1) model, for $\tau = \{0.01, 0.05, 0.1, 0.25, 0.5, 0.75, 0.9, 0.95, 0.99\}$. Numbers in brackets are bootstrapped p-values (2000 replications), calculated using the pair-wise bootstrap method for the $t(\tau)$ test. For the constant term we are testing the null of zero statistical significance, whereas for the slope coefficients we are testing the null of a unit root. One, two and three asterisks denote statistical significance at the 10, 5 and 1 percent level respectively.

Table 3.3 Autoregressive coefficients and estimated half lives

Panel A) Autoregressive coefficients and half lives (lag=1)

OLS	GBP	JPY	EUR
α	0.974 (∞)	0.981 (∞)	0.988 (∞)

Panel B) Quantile autoregressive coefficients and half lives (lags=1)

Quantile	GBP	JPY	EUR
	$a(\tau)$	$a(\tau)$	$a(\tau)$
1%	0.875 (5.191)	0.960 (∞)	0.992 (∞)
5%	0.958 (∞)	0.985 (∞)	0.988 (∞)
10%	0.964 (∞)	0.986 (∞)	0.988 (∞)
25%	0.987 (∞)	0.986 (∞)	0.994 (∞)
50%	0.988 (∞)	0.992 (∞)	0.993 (∞)
75%	1.009 (∞)	0.984 (∞)	0.992 (∞)
90%	0.984 (∞)	0.977 (29.78)	0.973 (∞)
95%	0.953 (14.398)	0.953 (14.398)	0.935 (10.313)
99%	0.898 (6.443)	0.912 (7.525)	0.918 (8.101)

(continued...)

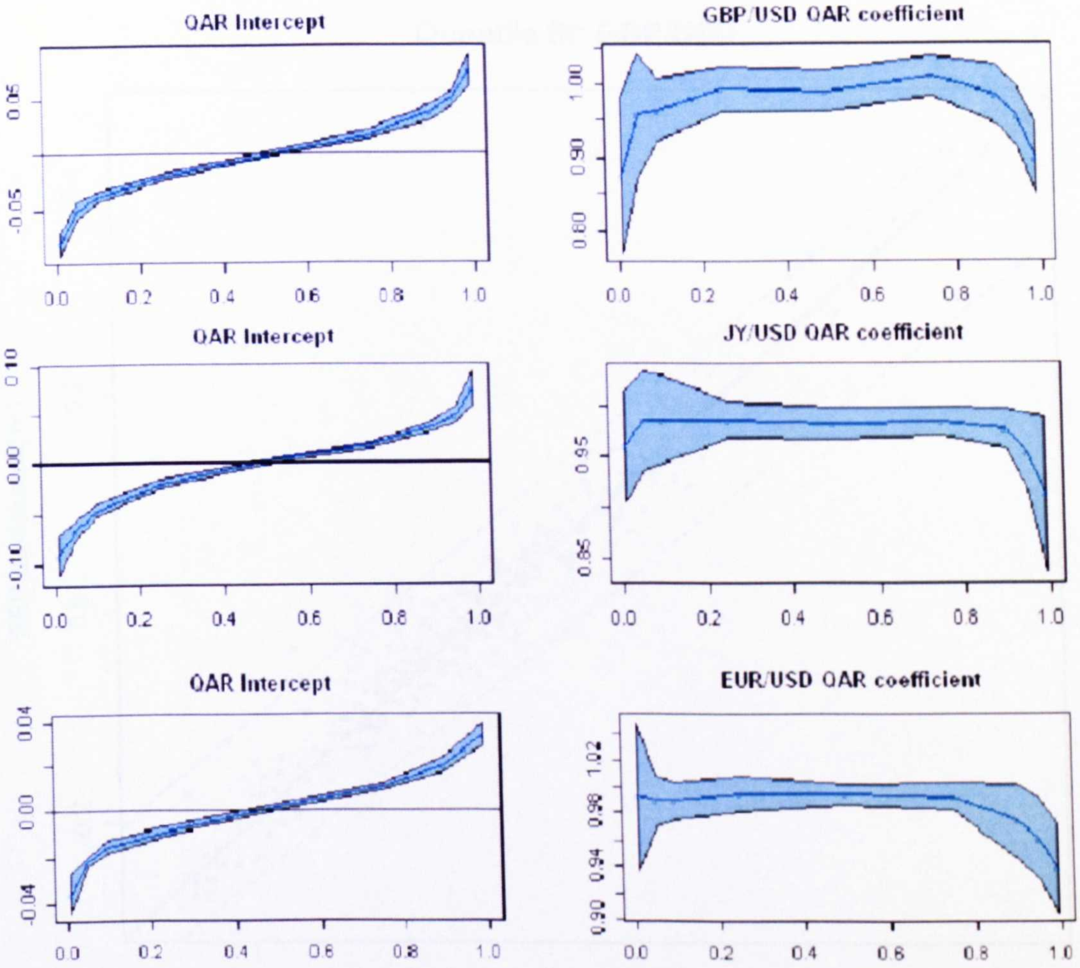
(...Table 3.3 continued)

Panel C) Non-parametric quantile autoregressive coefficients and half lives (lags=1)

	GBP			JPY			EUR		
Quantile	1%	50%	99%	1%	50%	99%	1%	50%	99%
$a(\tau)$	0.597 (1.344)	0.988 (∞)	0.789 (2.925)	1.050 (∞)	0.992 (∞)	0.855 (4.425)	0.925 (∞)	0.994 (∞)	0.719 (2.101)
	1.070 (∞)		1.004 (∞)	1.051 (∞)		0.950 (13.513)	0.954 (∞)		1.084 (∞)
	1.124 (∞)			0.898 (6.443)		1.006 (∞)	1.188 (∞)		0.761 (2.538)
	0.667 (1.712)			1.094 (∞)		0.937 (10.652)	1.053 (∞)		0.934 (10.152)
				1.004 (∞)		0.771 (2.665)	0.926 (9.016)		0.992 (∞)
				0.774 (2.706)			0.665 (1.699)		0.894 (6.186)

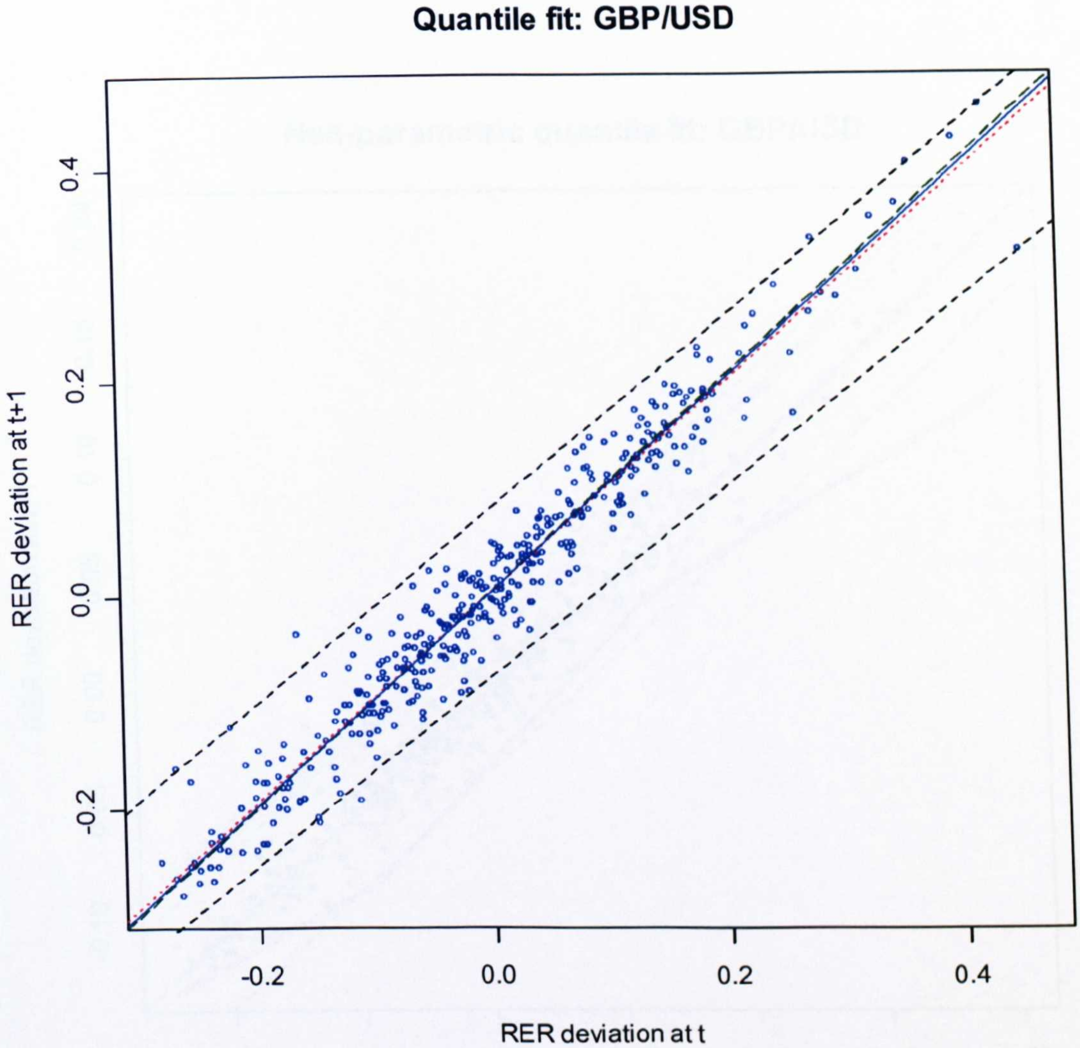
Notes: *Panel A)* The table presents the estimated values of the autoregressive (α) coefficient of a simple AR(1) model with the correspondent half lives in parenthesis, for each of the currencies under consideration. *Panel B)* The table shows the estimated values of the autoregressive ($a(\tau)$) coefficient of a QAR(1) model and the correspondent half lives in parenthesis, for $\tau = \{0.01, 0.05, 0.1, 0.25, 0.5, 0.75, 0.9, 0.95, 0.99\}$. *Panel C)* The table shows the estimated values of the autoregressive coefficient ($a(\tau)$), as they result from the non-parametric total variation penalty, quantile smoothing method, and their correspondent half lives in parenthesis for $\tau = \{0.01, 0.5, 0.99\}$. *For all panels* only the mean reverting coefficients (i.e. smaller than unity) were assigned a half live, whereas for the case of coefficients either bigger than unity or statistically not different from unity, half lives are set to infinity ∞. The significance of the coefficients with respect to unity, for the case of the non-parametric fit, was determined using asymptotic inference methods (He and Ng, 1999).

Figure 3.1 Quantile intercept and autoregressive (QAR) coefficients



Notes: The figures plot the quantile process of the intercept (right plots) and QAR coefficients (left plots) for each one of the major currencies. The vertical axis measures the values of the coefficients and the horizontal axis represents the values of the quantiles, ranging from *0.0* to *1.0*. The nine points on the plots are the coefficient (intercept and slope) estimates at $\tau=\{0.01, 0.05, 0.10, 0.25, 0.5, 0.75, 0.90, 0.95, 0.99\}$. The grey areas indicate the 95% confidence band.

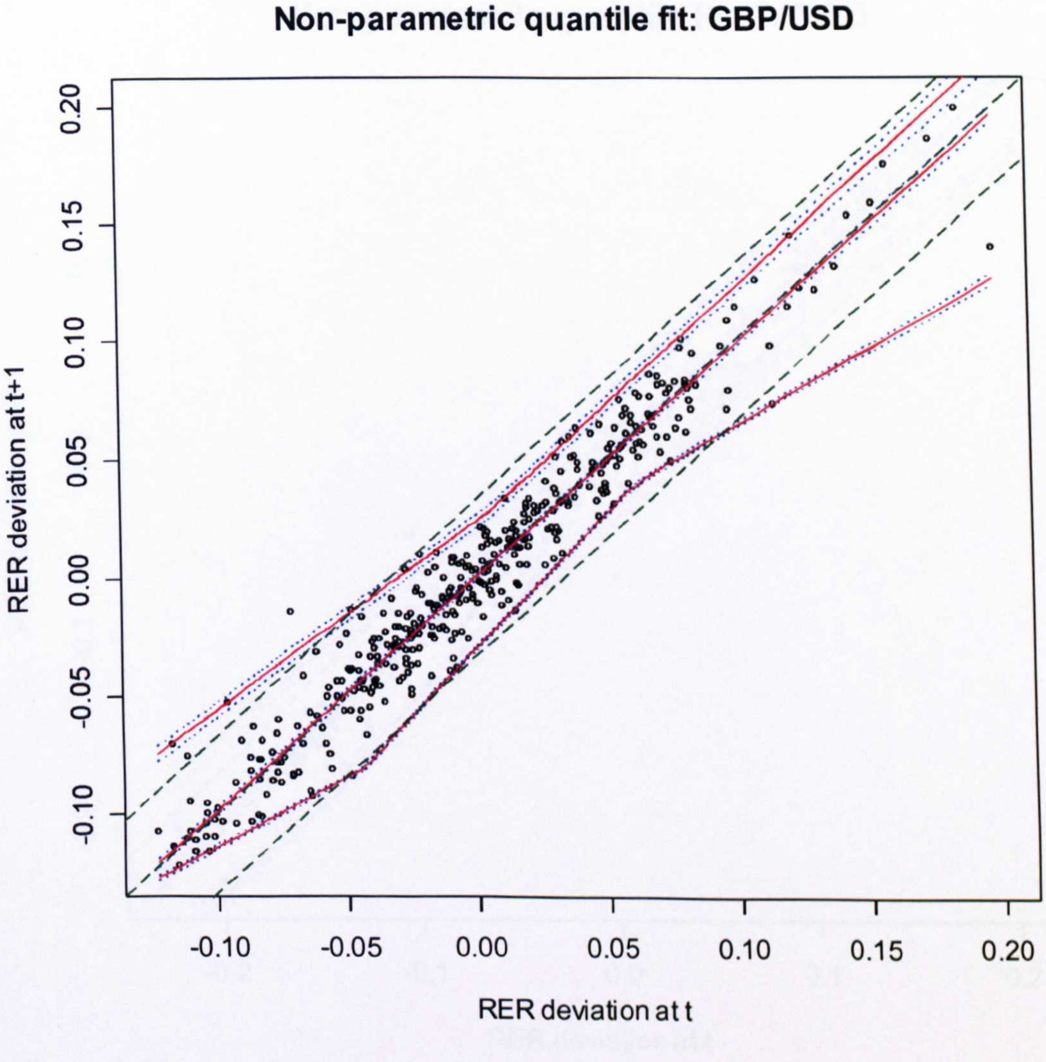
Figure 3.2 OLS and quantile fits



Notes: The figures present the realisations of the logged RER deviations (period *t-1* against period *t*) from Jan 1973 to Dec 2004 for the four currencies under consideration. The horizontal and vertical axes represent degrees of dis-equilibrium of the RER. The long-slash line represents the diagonal axis (x=y) and superimposed on that are the (OLS) mean and median fits, dotted and slashed lines respectively. The outer slashed lines represent the fit in the 1% (lower) and 99% (higher) quantile respectively. A detailed description is presented in Section 3.4.2.

Figure 3.3 Non-parametric quantile fit

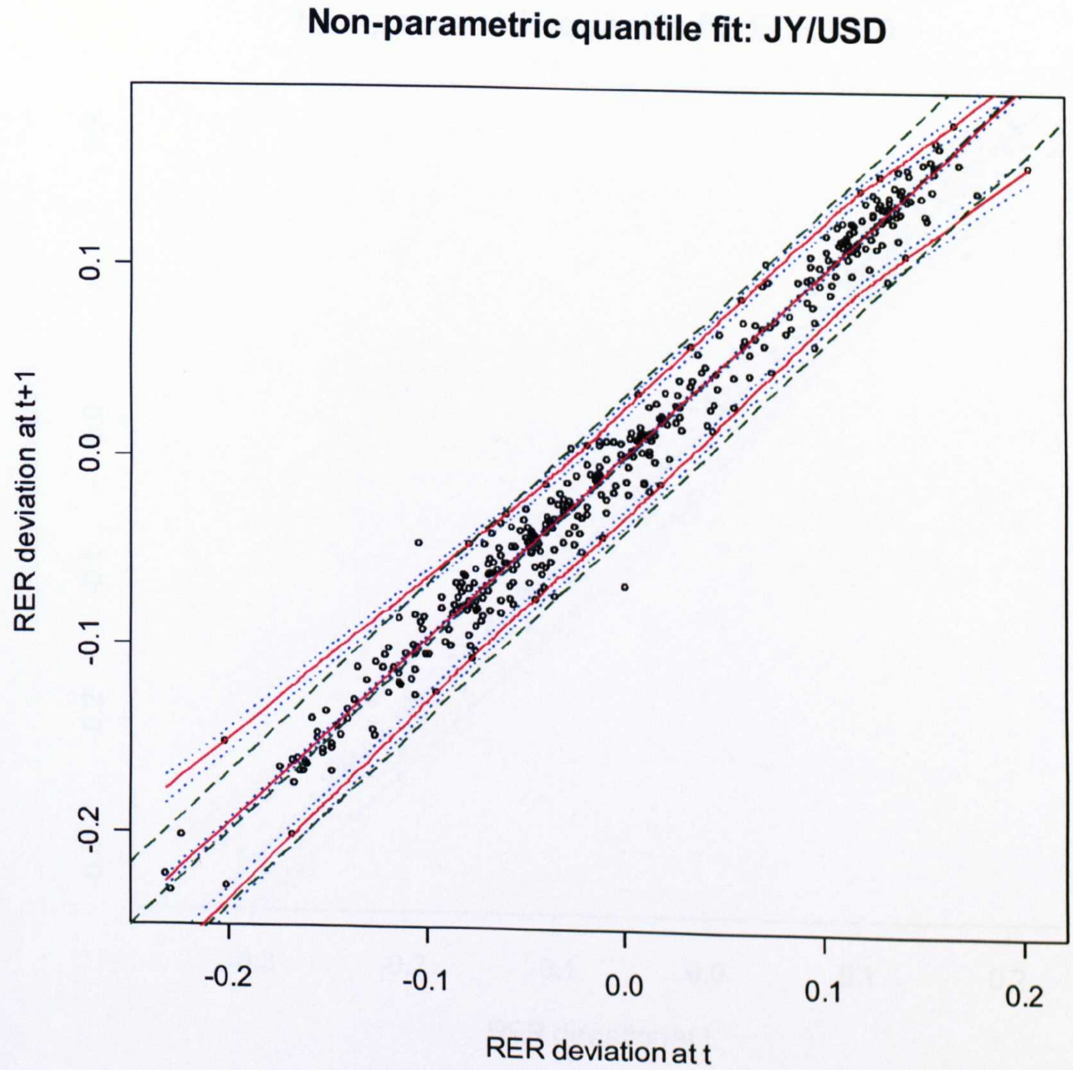
Panel A) British Pound



(continued…)

(…Figure 3.3 continued)

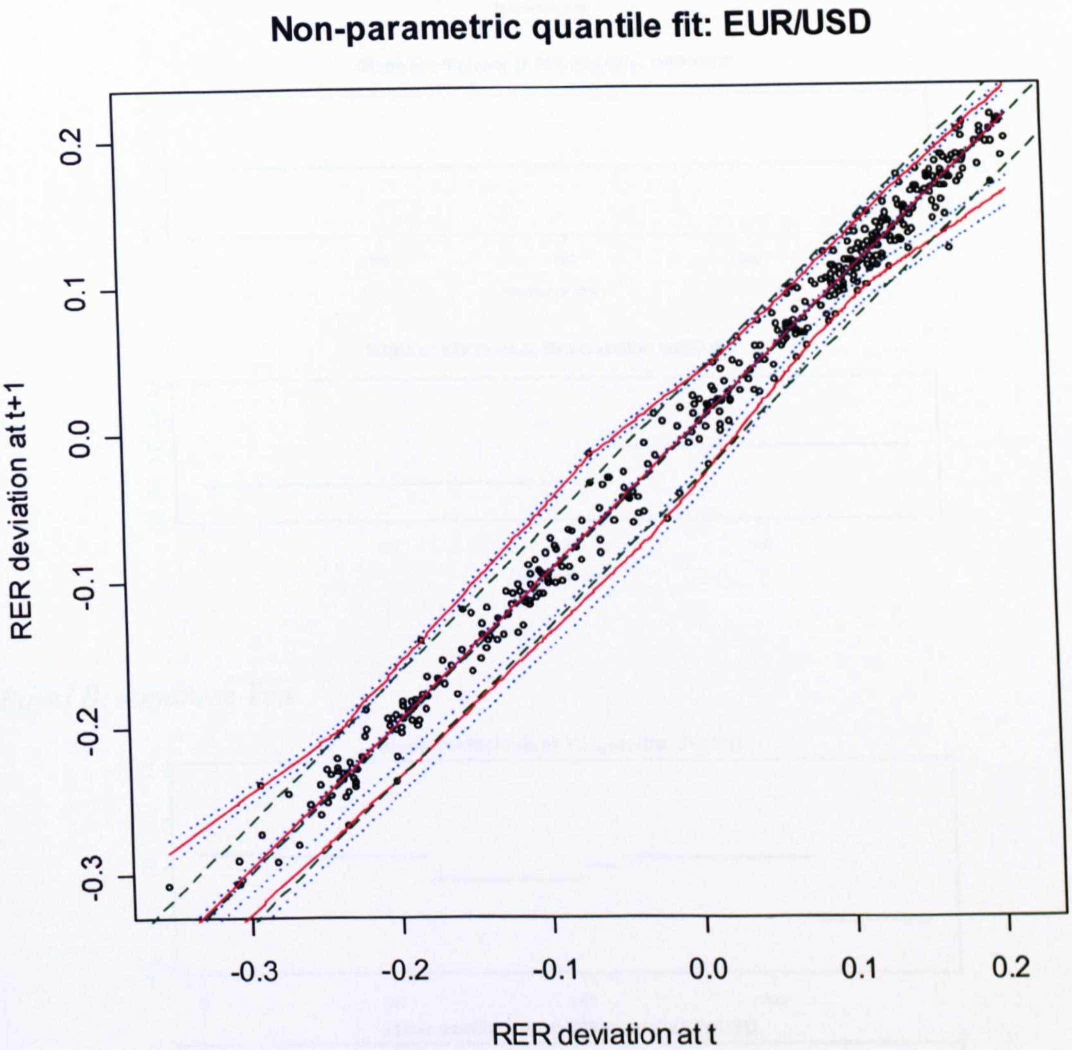
Panel B) Japanese Yen



(continued…)

(...Figure 3.3 continued)

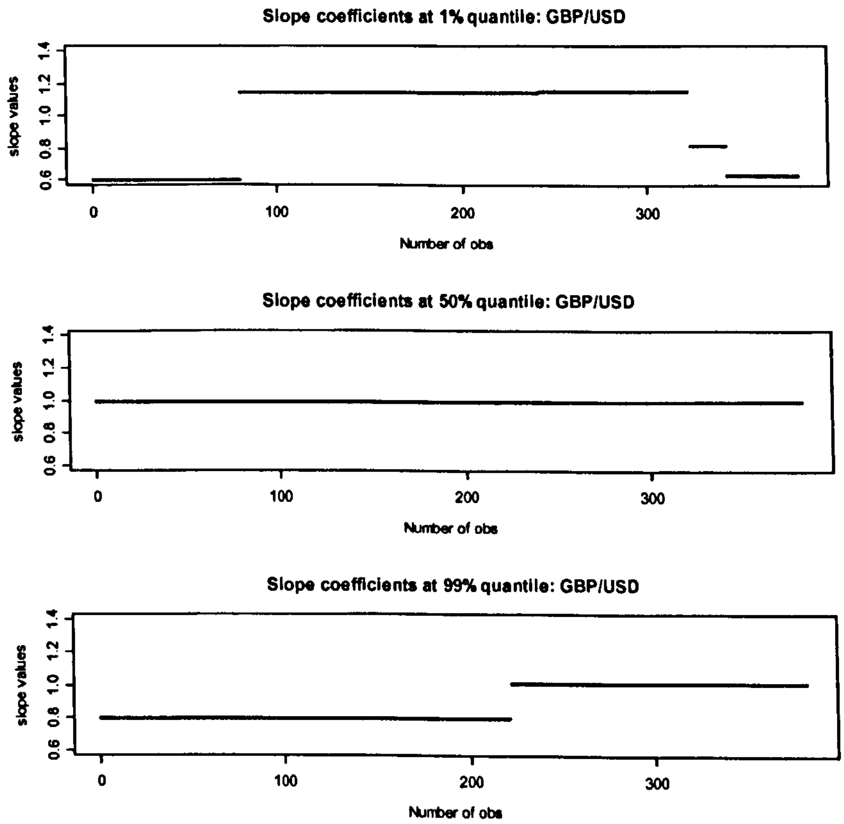
Panel C) Euro



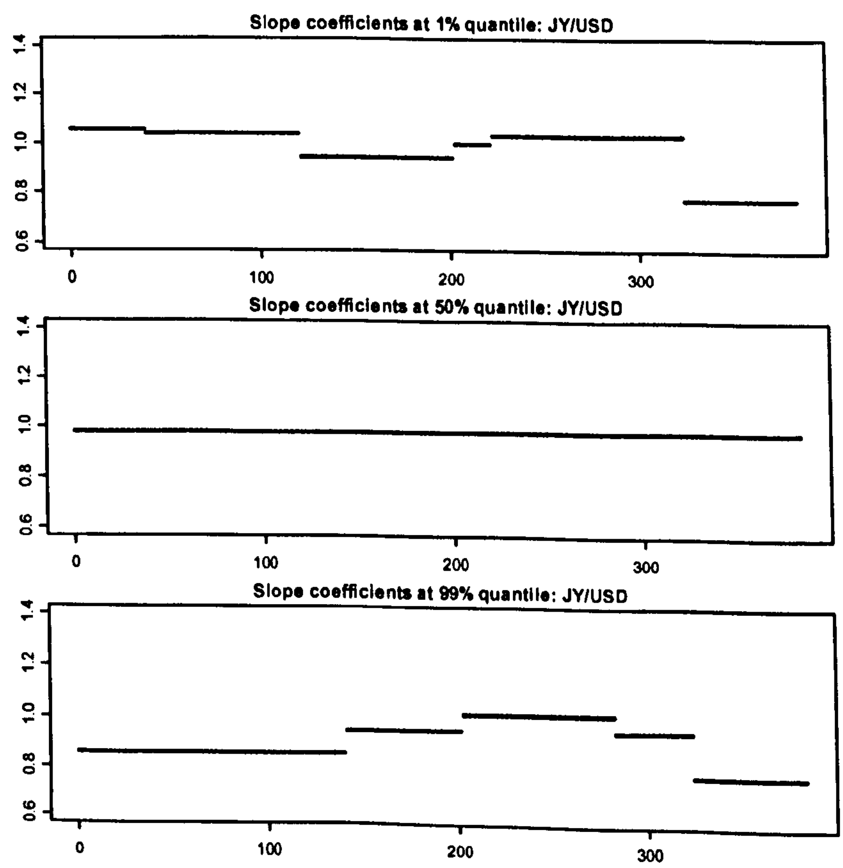
Notes: The figures present the realisations of the logged RER deviations (period *t-1* against period *t*) from Jan 1973 to Dec 2004 for the three currencies under investigation. The horizontal and vertical axes represent degrees of disequilibrium of the RER. Superimposed on the realisations are the fits of the regression quantiles smoothing splines for τ = *(0.01, 0.5, 0.99))* (solid lines), with standard error bands (dotted lines). The long-slash lines represent the diagonal axis (x=y) for the intercept values of the respective QAR fits (Table 3.2). *Panels A, B* and *C* refer to the analysis of the British Pound, the Japanese Yen and the Euro respectively.

Figure 3.4 Non-parametric quantile fit (slope coefficients)

Panel A: British pound



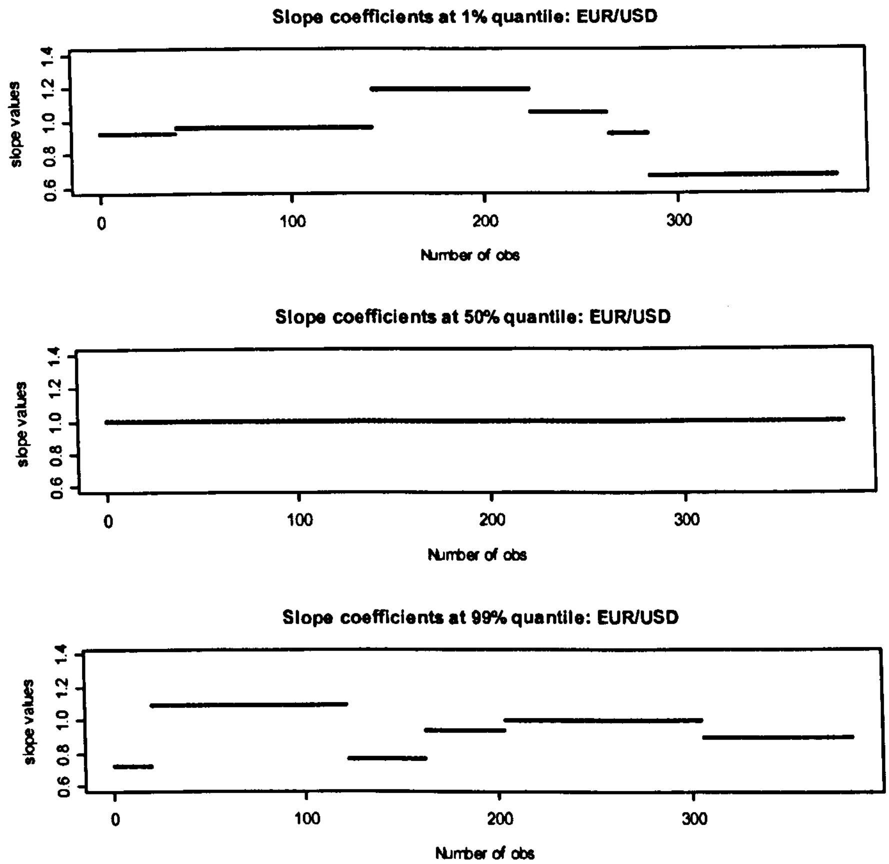
Panel B: Japanese Yen



(continued...)

(...Figure 3.4 continued)

Panel C: Euro



Notes: The figures present the correspondent slope coefficients at the different quantiles of the non-parametric quantile fits in Figure 3.3, for τ = *(0.01, 0.5, 0.99)*. The vertical axis presents the range of slope coefficient values and the horizontal axis the number of observations (relating to the ordered values of the RER realisations in the horizontal axis of Figure 3.3). *Panels A, B* and *C* refer to the analysis of the British Pound, the Japanese Yen and the Euro respectively.

Chapter Four

4 The relative importance of global versus domestic factors at driving money market interest rate differentials across countries

4.1 Introduction

The phenomenon of highly synchronised fluctuations in main economic variables across countries has captured a considerable amount of attention from academics and policy makers. More recently, the question on the effects of ever-intensifying globalisation and financial/economic integration, paired with advances in econometric techniques, have pushed the frontiers of research in this field and offered a renewed impetus in the exploration of common patterns in the fluctuations of major economic variables, with important policy implications. More specifically, it is recognised that when economic variables are driven by global forces, they may be less responsive to domestic policies, or influences by domestic causes. That would induce domestic authorities to adopt a more outward looking perspective for decision making (see e.g. Fisher, 2005; Borio and Filardo, 2006). More interestingly, if it becomes clear that certain countries have the status of a global player, thereby driving the behaviour of macroeconomic variables in other countries, policy responses of the rest of the world should be focused on the policies of these countries.

Along these lines a wide literature has provided evidence on the existence of strong comovements in business cycles, fostering the idea of a world business cycle (see Kose *et. al.*, 2004 and the references therein). At the same time evidence has also emerged of a world inflation rate (Mojon and Ciccareli, 2005). It is therefore expected that such strong international linkages in output and inflation should also induce strong comovements in monetary policies across countries.

The case of international monetary policy linkages becomes more interesting, if we accept that certain countries are considered as global "policy setters". In that case, it is possible to observe a certain degree of homogeneity or comovement in the response of the monetary policies of a group of countries vis-á-vis these global players, as this is expressed in money market differentials. Furthermore, it could be that there is a specific pattern in the responses of the rest of the world to the "policy setter", for example it could be that their common part is (implicitly) determined by the specific behaviour of one of those countries versus the "policy setter".

Close to this line of thinking, Frankel *et.al.* (2004) adopt a panel analysis to test whether the transmission of international interest rate changes to local rates is affected by the exchange rate regime. They find that over the last decade all exchange rate regimes exhibit high sensitivity of local interest rates to international ones. They also find that the US, Germany and Japan seem to be the only countries in their panel that can choose their own interest rates in the long run. Chinn and Frankel (2005) use a vector error correction model (VECM) to analyse the behaviour of world interest rates, considering both the US and Germany as global interest rate setters. They find that nominal US rates tend to drive European rates, although the former are becoming increasingly influenced by the latter. More recently, Diebold *et.al.* (2006) extend the Nelson-Siegel yield curve model to four countries and use dynamic factor analysis to extract the global yield factors (level, slope and curvature). They find evidence in favour of a global yield factor, which explains significant fractions of yield curve dynamics across countries. Finally, Dungey *et.al.* (2000) decompose international interest rate differentials of 10 year bonds into national and global factors. They use their results to rationalise international portfolio diversification and suggest the construction of an optimal portfolio, based on the global factor, which can outperform a simple equally weighted portfolio.

We use dynamic factor modelling and maximum likelihood estimation techniques in an effort to investigate the common fluctuations in the money market rate differentials of a group of major countries vis-á-vis a common denominating country. We follow the intuition of Frankel *et.al.* (2004) and Chinn and Frankel (2005) and use the US and Germany as the denominating countries. In our setting the interest rate differentials of each country are explained by a common/global factor and an orthogonal idiosyncratic, domestic component.

Our assumption is that the monetary policy stance of each country is reflected in the money-market interest rates, that is, interest rates with maturities up to one year. In that sense money market rates at different maturities reflect the different degrees of infiltration of monetary policy into the domestic money market rates at different horizons[42]. Therefore, the interest rate differential between two countries captures the divergence or deviation of the monetary policy stance of one country from that of the denominating country. By increasing the number of countries, we have a multiplicity of deviations, whose common parts, if at all existent, would be difficult to comprehend unless we can capture them by a common factor. Given that the countries represent the biggest economies in the world, we term this the global factor.

Overall, the global factor captures the common fluctuations in the monetary policy deviations of the countries under consideration with respect to the denominating country. More specifically, the global factor encapsulates the movements in inflation preferences, business cycles and further unaccounted risk premia that are shared within our group of countries but not with the denominating country. The inclusion of inflation

[42]This is not an unreasonable assumption, for credible and transparent central banks. In setting monetary policy, a central bank takes into account inflation considerations and the state of the economy. More specifically, central banks aim to achieve price stability in the short and long run and to stabilise the macro-economy in the short run. The degree of success depends on the credibility and transparency of the central bank. A credible and transparent central bank can successfully anchor inflation expectations (in the short, medium and perhaps long run, depending on the degree of credibility) and achieve economic stabilisation with their interest rate decisions. Thus risk premia due to uncertainties over inflation or output are, at least in the short run, negligible and only enter in the long run. Therefore it is possible, under these assumptions, to successfully steer short term interest rates, while effectively affect expectations about future interest rates.

tolerance and the state of the business cycle is straightforward, to the extent that these considerations are taken into account in monetary policy decisions. However, differences in the interest rate differentials might also be induced by other risk sources, due, perhaps, to movements in the exchange rates, political risk, global liquidity shocks, or the effects of oil prices. Such risk premia could drive a wedge in the interest rate differentials of two countries even when they share the same degree of inflation tolerance and are in the same phase of the business cycle. Therefore, the global factor also captures the part of the risk premium that is common across our group of countries vis-á-vis the denominating country.

This study is the first, to the best of the author's knowledge, to examine the existence, the nature and the implications of a common factor in short term interest rate differentials across countries. In that sense this study focuses not on the similarities but on the discrepancies between monetary policies of a group of countries vis-á-vis the US (German) policies. By exploring the extent to which these discrepancies are driven by a common force, we provide further evidence of integration in a globalised environment, with direct implications for policy makers.

More specifically, we reveal valuable insights in the behaviour of the short-term interest rate differentials with direct implications for monetary policy actions. First, we get a measure of interest rate differentials comovements across countries. Second, we get an indication of countries with global status in the sense that they affect the behaviour of the global factor. To this direction, our methodology allows us the advantage to investigate finer interactions among global policy makers. Namely, we can see whether the common part of policy divergencies vis-á-vis the denominating country is driven by the behaviour of a specific country. Third, we further investigate whether global players can act as interest rate setters, that is, set their interest rates independently. Fourth, we draw inferences on the existence of a world interest rate.

We report evidence of strong comovement in the interest rate differentials of a group of countries vis-á-vis the denominating country. In other words, there seems to be a strong global factor driving the interest rate differentials of several countries, suggesting increased sensitivity of the domestic interest rate differentials to the global one. However, it is notable that sensitivity patterns might change depending on the denominating country. Nevertheless, the global factor seems to be reacting to monetary policies of the US and the EU. In fact, it appears that when the US is the denominating country, the global factor seems to be driven primarily by EU policies vis-á-vis the US. Nevertheless, the Euro-area and the US seem to enjoy a certain degree of independence when choosing their own policies. Finally, there seems to be evidence that the US policy rate emerges as a global interest rate.

In terms of monetary policy implications, our results suggest that although there are still discrepancies between the monetary policies of various countries, these discrepancies are not erratic, but rather coordinated and influenced by major global players like the EU and the US. Therefore, monetary policy-makers should pay closer attention to foreign macroeconomic aggregates and, therefore, foreign monetary policy choices of these global players.

Our analysis proceeds as follows. Section 4.2 demonstrates the econometric methodologies used throughout this paper. Section 4.3 describes the data sources and the construction of our final data set. Section 4.4 presents, evaluates and analyses the main results and the results from the robustness analysis and, finally, Section 4.5 concludes.

4.2 Methodology

Our methodology involves two steps. In the first step we use dynamic factor analysis and maximum likelihood estimation techniques to extract the common fluctuations across the interest rate differentials of interest. We also measure the relative contribution of the common factor to variations in interest rate differentials for each country. In the second step, we investigate the relationship between the common factor and the monetary policies of the US and EU, in an effort to identify the driving forces of the global factor.

4.2.1 Dynamic factor analysis

The main idea is to decompose the observed series of interest rate differentials into an unobserved common component and a country specific component. The econometric model used to achieve that follows the literature of the unobserved component, dynamic factor models, proposed by Sargent and Sims (1977) and Geweke (1977) and further studied by, amongst others, Harvey (1989) and Stock and Watson (1991). These models capture the common features (correlations) amongst economic series in a single factor, the common factor. In contrast to static factor models (i.e. principal components), dynamic factor models allow the underlying factor to evolve dynamically, so they have the advantage of measuring contemporaneous and temporal comovements among the variables.

The model considered is a simple single factor, state-space model, where the interest rate differentials are represented by the sum of a global component, i.e. a factor common to all countries, and an orthogonal, country specific, idiosyncratic component. More precisely, assume a vector of size N, which contains the observable interest rate differentials of maturity m for country c, where c=1 to N, vis-á-vis the denominating country at time t, denoted Y_t^m. In our case $N = 7$ and $m = 3, 6, 9$, and 12-months. Assume F_t^m is a single measure of unobserved common fluctuations in the interest rate differential series, i.e. the common factor. ε_t^m is the country-specific, idiosyncratic component. Then the measurement equation of our state space model can be written as:

$$Y_t^m = A'F_t^m + \varepsilon_t^m, \tag{4.1}$$

or, more analytically

$$\begin{bmatrix} Y_{t,1}^m \\ \cdot \\ \cdot \\ \cdot \\ Y_{t,N}^m \end{bmatrix} = \begin{bmatrix} A_1 \\ \cdot \\ \cdot \\ \cdot \\ A_N \end{bmatrix} * \begin{bmatrix} F_t^m \end{bmatrix} + \begin{bmatrix} \varepsilon_{t,1}^m \\ \cdot \\ \cdot \\ \cdot \\ \varepsilon_{t,N}^m \end{bmatrix}, \tag{4.2}$$

where A is a size N vector containing the coefficients or factor loadings for each country. That is, each coefficient in the A vector measures the effect of the common factor on the respective interest rate differential series. The idiosyncratic component is assumed to be normally distributed, with zero mean and constant variance $E(\varepsilon_t^m \varepsilon_t^{m\prime}) = \Sigma_v$, where Σ_v is a $(N \times N)$ diagonal matrix. In this model the common factor is uncorrelated with the error, $E(F_t^m \; \varepsilon_t^{m\prime}) = 0$. That is, the common and the idiosyncratic, country specific components are orthogonal, therefore the decomposition between them is exact.

The common factor is dynamic and is modelled as a first-order autoregressive or AR(1) process. The transition equation is as follows:

$$F_{t+1}^m = \rho_F F_t^m + w_t, \tag{4.3}$$

where ρ_F is the autoregressive coefficient measuring the persistence of the series. Note that we assume that the dynamic factor follows a stationary process, given that interest rate differentials amongst countries are stationary processes (see Dungey *et.al.*, 2000; Sarno *et.al.*, 2007 and references therein). The innovation w_t is assumed to be zero mean, contemporaneously uncorrelated normal random variable, orthogonal to the global factor, $E(F_t^m \; w_t^\prime) = 0$. The estimation of the unobserved factor is carried out by maximum likelihood estimation[43].

Estimation of the unobserved components allows us to measure the relative contributions of the common versus the idiosyncratic factor for each country's interest rate

[43] Alternative methods based on principal components and Kalman filtering (see Stock and Watson, 1991) can and have also been used in this setting as robustness checks and yield similar results.

differential. We do that by computing the share of the variance of each variable Y^m due to the global factor. Given that the global and the domestic factor are orthogonal, the decomposition is exact and therefore,

$$var(Y_t^m) = A'var(F_t^m)A + var(\varepsilon_t^m) \tag{4.4}$$

Given $var(F_t^m)$=1, $var(A'F_t^m) = A'A$. In that case, the percentage of variance in each country c explained by the common factor can be calculated as $A_c^2/var(Y_{t,c}^m) * 100$ and by the domestic factor as $var(\varepsilon_{t,c}^m)/var(Y_{t,c}^m) * 100$ and

$$A_c^2/var(Y_{t,c}^m) + var(\varepsilon_{t,c}^m)/var(Y_{t,c}^m) = 1. \tag{4.5}$$

Therefore, the relative importance of the global versus the domestic factor could be seen by checking the percentage contribution of the global factor for each country c, $A_c^2/var(Y_{t,c}^m)$. The closer the number is to one, the bigger the importance of the global factor in explaining the variance of the series relative to the domestic factor.

In our context, a statistically significant global factor that explains a large part of the variance of the interest rate differential would suggest that the country under consideration has a high degree of harmonisation/integration with the rest of the countries in its response vis-á-vis the denominating country. On the contrary, a statistically insignificant global factor, or one that explains a small part of the variance of the interest rate differential (low R-squared) would suggest that the domestic factor is the one driving the interest rate differential. Therefore, this country's monetary policy stance is different vis-á-vis the denominating country for reasons that are idiosyncratic to that country and are not shared by the rest of the countries.

Finally, this analysis can also signal the relative importance of the denominating country. A strong global player is more likely to attract a common interest to its policies. Therefore, the stronger the global factor (in both statistical and economic terms), the more homogenous and coordinated would be the comovement of the rest of the world to the monetary policy stance of the denominating country.

4.2.2 The common factor and monetary policy stance

Following the estimation of the global factor, we then try to rationalise its driving forces. In the existing literature a number of fundamentals has been posited as possible explanators of the interest rate differentials. These include variables that relate to economic integration, such as capacity utilisation and output levels, ratios of debt, current account balances, international reserves, investment to GDP ratios[44] etc.. Other potentials explanators relate to financial integration, such as the openness and functioning of financial markets globally, linked to the imposition of controls or restrictions to financial instruments. These variables combined determine country specific risk premia and the reaction of each country to a common stage of the business cycle. Thereby, they affect monetary policy decisions and exchange rate risk.

We want to measure the extent to which monetary policies and exchange rates of major global players affect the policy stance of the global economy. This is done by regressing the global factor on monetary policy rates, other policy variables and exchange rate changes. The setting is ideal, since the global factor solely captures the part of the monetary policy response that is due to external influences. We focus on the global factor with respect to the US and consider as regressors policy variables of the EU and the US separately (EU or US) and combined (EU/US). This setting allows us to assess the international role of these two countries as global players and interest rate setters, in line with the recent strand of literature on the global interest rate (see Chin and Frankel; 2005, Frankel *et. al.*; 2004). We use two types of models

$$M_{1,c} : F_t^m = \alpha_c + \beta_{1,c} r_{t-1,c} + \gamma_{1,c} FX_t^m + \zeta_{t,c} \tag{4.6}$$

[44] Arguably, debt, current account balances and investment decisions mainly affect long term interest rates, whereas short term rates relate more to output gaps (see for example Ang and Piazzesi, 2000).

and

$$M_{2,c}: F_t^m = \kappa_c + \beta_{2,c} r_{t-1,c} + \delta_c inflation_{t+1,c} + \lambda_c output\_gap_{t+1,c} + \gamma_{2,c} FX_t^m + \eta_{t,c} \quad (4.7)$$

where $M_{1,c}$ and $M_{2,c}$ stand for Model 1 and Model 2 respectively for country c, where c now is restricted to c=US, EU, US/EU. $M_{1,c}$ uses as explanatory variables $r_{t,c}$ and FX_t^m, which represent the overnight (policy) rates -Eonia for the EU and Federal Funds (FF) for the US- and the common factor of expected exchange rate changes respectively[45]. In the second model, $M_{2,c}$, we also allow the common factor to be explained by other relevant macroeconomic variables. We, therefore, include $inflation_{t+1,c}$ and $output\_gap_{t-1,c}$, which represent inflation and the output gap respectively. α_c and κ_c are constants, and $\beta_{i,c}, \gamma_{i,c}, \delta_c, \lambda_c$ are the coefficients of the policy rate, the exchange rate, inflation and the output gap respectively for $i = 1, 2$, referring to the first ($M_{1,c}$) or the second ($M_{2,c}$) model respectively.

Notice that the policy rates in those equations are included with a lag. These interest rates directly relate to the monetary policy stance, and should therefore take into account interest-rate smoothing considerations, hence the use of the lag. In our setup, estimates of these regressions could signal the extent to which monetary policy decisions in those two countries drive the global factor of money market rates. We also include one-period leads for inflation and the output gap as well as the common factor of the expected exchange rate change. These variables reflect further inflation, output and exchange risk premia considerations that affect the global factor and are not already captured by the monetary policy rates. Finally, the constants should capture various other unaccounted risk premia.

The models are estimated by ordinary least squares (OLS), which assumes that the monetary policy variables of the US and the EU are exogenously determined. More

[45] For further information on the construction of FX_t^m see Section 4.3.

intuitively, the US and EU monetary authorities are setting their monetary policies without considering the common reaction of the rest of the countries; in other words, they are able to choose their own interest rates. In that case, the common factor is affected but does not affect the monetary policy stance of these two countries.

Failure to validate this assumption would introduce endogeneity problems in the regression, which would render the above OLS estimates inconsistent. For this reason we re-estimate the models presented above using instrumental variables (two stage least squares) estimation techniques as a robustness check. Lagged variables of the dependent variables are being used as instruments. The validity of the instruments is being tested with Sargan's test of overidentifying restrictions[46]. Under the null hypothesis, all instruments are uncorrelated with the error term. Lack of endogeneity would suggest that the monetary policies of these countries are indeed exogenously determined and therefore independent.

Overall, the above simple regressions shed light on the following issues. First, we measure the extent to which monetary policies and exchange rates of major global players drive the policy reactions of the global economy. More specifically we can identify whether policy divergencies between the EU and the US are being followed by the rest of the countries. Second, we see the relative importance of US versus EU policies in driving the factor and therefore confirm whether the US rate can be considered as the world interest rate. Third, we can judge the role of the US (EU) as independent interest rate setters. Finally, it is an implicit way of testing our original assumption, namely that policy rates are reflected in the short-term, money-market rates.

[46] A potential instrument should be correlated with the dependent variable, in our case the common factor, but not with the error term. In practice, a sufficient condition of instrument validity is that the covariance matrix of the instruments has full rank. Sargan's test for overidentifying restrictions tests the joint null hypothesis that the model is correctly specified and that the instruments used are valid.

4.3 Data

Our data set consists of interest rate (bond yields) series, exchange rate variables and macro variables. Our sources include Ecowin (Reuters) and Bloomberg for the interest rate variables, and the International Monetary Fund (IMF)'s International Financial Statistics (IFS) and the Organisations of Economic Cooperation and Development (OECD)'s, Main Economic Indicators (MEI) for the exchange rate and macro variables.

The interest rate series relate to monthly bond yields with maturities of 3, 6, 9 and 12 months for a set of eight countries (Canada, France, Germany, Japan, Norway, UK, US) spanning a common sample from January 1985 to March 2006. In order to obtain a series for each maturity and each country, we have used smoothing spline extrapolation, whereby a cubic spline is fit to the observed data points (see Waggoner, 1997 and Anderson and Sleath, 2001 for applications of the smoothing spline method to the term structure of interest rates).

Exchange rate variables vis-á-vis the US dollar and the synthetic DM/euro were used to extract the common (global) component of the expected exchange rate change. This variable is similar in construction and spirit to the global component of the interest rate differentials. Namely, a similar state space model is used, where the expected change in the exchange rate of countries vis-á-vis US and Germany for horizons of 3, 6 and 9 months ahead is driven by an unobserved factor, common to all exchange rates and an orthogonal, idiosyncratic factor with the same properties as in the model of section 4.2.1[47].

Turning our attention to the macro variables, our data set includes monthly observations of policy rates, inflation and the output gap for the US and the EU. The policy rates are taken to be overnight rates from the inter-bank market for the US and the EU,

[47] Justification for this approach can be found in Dungey *et.al.* (2000), who appeal to Uncovered Interest Parity (UIP) to suggest that both foreign exchange returns and interest rate differentials share the same factor model.

that is the federal funds rate and the eonia rate respectively[48]. Interest rate data are not logged. Inflation for the US and Germany is constructed as the twelfth difference of the logged CPI values of the US and the EU. The output gap for the two countries is extracted by applying a Hodrick-Prescott filter on the log of the real industrial production for each country respectively.

4.4 Results

4.4.1 Global versus domestic factors

This part of our analysis presents and attempts to interpret the results from the state space model described in Section 4.2.1. Table 4.1 shows the main results concerning the global factor contributions. Starting with Panel A, the factor loadings suggest that the global factor is in every case statistically significant and positive. Also, it would be relevant to notice that the loading coefficients are bigger in the case of the US as the denominating country, than in the case of Germany, suggesting that the weight of the global factor is bigger in explaining interest rate differentials with respect to the US rather than Germany.

However, more robust insights on the relative importance of the global versus the domestic factor can be drawn by looking at Panel B, which presents the percentage of variance explained by the global factor with respect to the US and Germany. An overall look suggests that the global factor is quite dominant for both denominating countries and explains, in the vast majority of cases, more than 50% of the variance of the interest rate differentials. In other words, it seems that there is a high degree of homogeneity in the monetary policy divergencies of our group of countries vis-á-vis the monetary policy choices of the denominating country.

[48]Note that the interbank market is different from the money market both from an economic and institutional point of view. For a description of the institutional and economic characteristics of the interbank market see Edwards (1986).

Nevertheless, a closer look can reveal more subtle differences in the importance of the global factor both within and across the denominating countries. Focusing our attention to the results of the US as denominating country, we can see that in all cases the global factor explains at least half of the variance of a country's interest rate differentials. The highest percentage explained is in the case of the German differential (95%). Japan and Switzerland record numbers not far behind. Overall, these results are not implausible, given that these countries are important players, and we would therefore expect a strong interaction amongst them vis-á-vis the US. The other EU countries, the UK and France, appear less integrated in their behaviour, which seems to be partly (50%) driven by intrinsic forces.

Turning our attention to the global factor with respect to Germany, we can observe that the US differential displays a similar behaviour to the German differential. In particular, the global factor explains almost 85% of the US differential for almost all maturities. However, the global factor also explains most of France's interest rate differentials, and in fact marginally more than the US's, with the UK following rather closely. This result could reflect the fruit of the European integration process, or simply stress the importance of the geopolitical vicinity and ties of these countries with respect to Germany. However, the global factor seems to have a much smaller involvement in determining the policy deviations of Switzerland and Japan. Notably, the case of Japan suggests the policy deviations of Japan versus Germany are quite idosyncratic and clearly not shared by the rest of the countries in our sample.

For both cases, Norway seems to be only partly (50%) driven by global and domestic factors vis-á-vis both the US and Germany. This is a logical outcome if we consider the peculiarities of the Norwegian economy (the only major oil-producing country in our sample), which makes it subject to different business cycles, therefore adjusting its policy to idiosyncratic shocks. Finally, for the case of Canada the global factor with respect to Germany seems to be explaining more of the interest differential variation than the global factor with respect to the US, despite Canada's vicinity and links to the US economy. Such results find support in the study of Dungey *et.al.* (2000)[49]. In any case, the difference is only marginal and in both cases the percentage of variance explained is quite high, suggesting the harmonisation of Canadian policies with global practices.

Overall, the above analysis suggests that the global factor explains a big part of the interest rate differential variance for most countries, thereby policy deviations of the various countries with respect to the policy stance of the US or Germany are broadly harmonised and globally integrated. Furthermore, the global factor explains almost all variation in the US and German differentials. Taken together these findings could perhaps hint a dominant role of these two countries in global interactions. In order to further investigate this notion, and taking into account the literature that treats the US interest rate as the world interest rate, we will try to identify the driving forces of the global factor by appealing to US and EU monetary policies. In the following section we present the results from regressing the global factor on various monetary policy instruments of the US and the EU, in order to measure the extent to which these two areas affect the global factor.

4.4.2 Monetary policy and the global factor

Main analysis: As a first step, we present the contemporaneous correlation coefficients of the two global factors (with respect to the US and Germany) with the money market interest rates of the US and Germany for each maturity (Table 4.2). The results show that the global factor, be it with respect to the US or with respect to Germany, appears to be strongly, positively correlated with the US interest rate across all maturities.

[49]Further support towards this direction can also be provided by IMF data sources, such as the Coordinated Portfolio Investment Survey (CPIS). Looking at the bilateral portfolio investment assets plus liabilities between the United States and each partner country, one could observe that financial linkages between the US and Canada are not as pronounced as one might wish to entertain.

From Table 4.2 two things become clear. First, the interest rates of the denominating country do not necessarily drive the behaviour of the global factor. Second, the global factor could, to a certain extent, be driven by the US interest rates. This conjecture bodes well with the literature which terms the US rate as the world interest rate. However, a more formal and robust investigation is warranted, therefore we estimate the relative weights of US and EU monetary policy variables on the global factor. As already mentioned in Section 4.2.2, we focus our analysis on the global factor with respect to the US.

Table 4.3 presents the results of the regressions described in Section 4.2.2. Looking at the different model specifications, what becomes directly observable is the prevalence of the policy rates in explaining the global factor, as compared to other policy variables. Namely, policy rates are not only statistically significant, but their inclusion significantly increases the R-squared measure of overall fit. When accounted on its own, the federal funds rate appears to be an important explanatory source of the global factor, judging from the R-squared measure. This further enhances evidence on its role as the world interest rate. However, the inclusion of both policy rates clearly produces the best fit, with the relative weight of the two policy rates (absolute value of the estimated coefficient) being almost the same (the federal funds rate is only marginally higher). This result clearly suggests that the interaction of US and EU monetary policies drives the global factor. More intuitively, the policy divergencies of our group of countries vis-á-vis the US mainly follows the pattern set out by the EU vis-á-vis the US.

With regard to the rest of our explanatory variables, the inflation and output gap appear to be insignificant in almost all cases, more so at shorter maturities. This could suggest that further inflation and output considerations potentially affecting the global factor are already captured by the monetary policy rates. However, the strong significance of the term that represents the expected exchange rate changes and the (sometimes significant) constant term suggest that other risk factors, apart from the ones included in the policy rates might drive the global factor. More precisely, the exchange rate factor is always strongly significant and seems to be explaining an important part of the global factor[50]. Nevertheless, this result suggests that common perceptions about future appreciations or depreciations could drive similar responses of monetary policy actions. Lastly, the significance of the constant term suggests the existence of further, unidentified risk premia.

Overall, our results demonstrate a robust sensitivity of the global factor to the interaction of the US and EU policies and exchange rate risk. Judging from the relevant contributions to the overall fit, the model best describing the behaviour of the global factor combines both EU and US policy rates and the exchange rate factor, achieving an overall fit of more than 80% for almost all maturities. The largest part of this result is driven by the inclusion of both policy rates (more so the US ones) whereas a smaller part is owed to foreign exchange related risk. In that respect our model supports the international role of the US and the EU policies in affecting policy responses of the rest of the countries. Also, our results highlights the role of the US rate as the global interest rate.

Robustness check: Finally, in order to account for any kind of endogeneity, we repeat the previous exercise using two stage least squares estimation and report the results in Table 4.4. The main conclusion is that the results do not change qualitatively, and the quantitative differences are small. We only report results for the 3-month horizon due to space preservation considerations. However, the conclusions drawn apply to the rest of cases as well. More specifically, comparing the estimated coefficients between Table 4.4 and Table 4.3 (Panel A) we find little quantitative differences. Given that our instruments are valid, as suggested by the J-statistic, our results imply that the OLS

[50]It seems from our experiment that the interest rate differential factor inherits part of the factor movements in the exchange rate movements. Although we do not purport to present evidence in favor of UIP, we provide this result as an interesting finding which could be investigated in further UIP research. We should, however note that, in our data set, the relationship between the interest rate differential and the expected exchange rate change breaks down in the usual two country setting.

regressions did not suffer from endogeneity and therefore the estimations in Table 4.3 are robust. By implication the interest rate differential between the US and EU affects the global factor, but is not affected by it. Therefore these two countries can set their policies exogenously and independently.

4.5 Conclusion

In this paper we aim to analyse the international linkages in the money market rate differentials with respect to the US and Germany across a group of countries. We interpret these differentials as reflecting the monetary policy deviations of each country versus the denominating country (US or Germany). We employ a maximum likelihood, dynamic factor model where the interest rate differentials of a group of countries vis-á-vis a common denominating country are modeled as the sum of a common/global component and an idiosyncratic/domestic component. In that sense, the global factor captures the common fluctuations in the policy responses of our group of countries versus the denominating country. We provide an analysis of the relative importance of global versus domestic factors and we also try to identify the special role of the two major economies, the US and the EU in driving the global factor itself.

The main contribution of our study relies on the use of factor analysis methods on short term interest rate differentials. Adding to previous studies, our study provides a measure of comovements in the monetary policy reactions across countries and the importance of these comovements for the policy fluctuations of each country under investigation. We also get an indication of the existence of dominant countries, either in the sense that they affect the strength of the global factor or that their policy fluctuations drive the global factor fluctuations. We then further investigate finer interactions amongst our players, and specifically whether the policy deviations of a group of countries from the denominating country are explained by the pattern set by the deviations of another dominant country. We also explore whether these dominant countries can set their policies independently of others and yet command other countries' responses. Finally, we provide indications for the role of the US as the global interest rate.

We reveal valuable insights in the behaviour of the short-term interest rate differentials with direct implications for monetary policy actions. We find that there is a significant common global component present in the fluctuations of the interest rate differentials for almost all countries, although the patterns of interaction seem to be different for the US and Germany, revealing different spheres of US and German policy influence and, potentially, different degrees of regional integration. For example, EU country differentials are much more integrated vis-á-vis Germany than vis-á-vis the US. The opposite is happening for Japan and Switzerland. Based on this result, we can further argue that the US and the EU are in a position to choose their policies independently, although the interaction of their policies seems to be followed by the rest of the world. Finally, our analysis suggests that the US interest rate is the dominant force behind the global factor, thereby providing support for the role of the US as the global interest rate.

Overall, our analysis accepts that the monetary policies between a set of countries may be different. However, we find that a large part of these differences are in fact responding to common considerations and are shared across countries. More importantly, they tend to follow the policy interactions of two major global players, namely the US and the EU. Therefore, the policy implication would be that monetary policy-makers should pay closer attention to US and EU policies when forming their own.

Table 4.1 Global factor contributions

Panel A) Global factor loadings

	Common factor wrt US					Common factor wrt Germany			
	3-month	6-month	9-month	12-month		3-month	6-month	9-month	12-month
GE	0.977 (0.021)	0.978 (0.020)	0.978 (0.020)	0.983 (0.018)	US	0.697 (0.048)	0.715 (0.047)	0.884 (0.033)	0.694 (0.049)
UK	0.747 (0.060)	0.759 (0.059)	0.796 (0.055)	0.771 (0.058)	UK	0.506 (0.063)	0.503 (0.063)	0.681 (0.052)	0.639 (0.053)
CA	0.802 (0.054)	0.805 (0.054)	0.751 (0.060)	0.809 (0.054)	CA	0.469 (0.065)	0.491 (0.064)	0.477 (0.066)	0.561 (0.059)
FR	0.715 (0.064)	0.723 (0.063)	0.509 (0.078)	0.952 (0.029)	FR	0.714 (0.046)	0.726 (0.046)	0.893 (0.032)	0.271 (0.077)
JA	0.919 (0.036)	0.918 (0.037)	0.918 (0.037)	0.892 (0.041)	JA	0.079 (0.087)	0.084 (0.087)	0.070 (0.087)	0.103 (0.086)
NO	0.731 (0.062)	0.767 (0.058)	0.799 (0.058)	0.818 (0.052)	NO	0.331 (0.073)	0.351 (0.072)	0.283 (0.077)	0.322 (0.074)
SW	0.947 (0.030)	0.952 (0.029)	0.942 (0.029)	0.943 (0.031)	SW	0.267 (0.077)	0.246 (0.078)	0.124 (0.085)	0.064 (0.088)

Panel B) % of Var explained by Global factor

	Common factor wrt US					Common factor wrt Germany			
	3-month	6-month	9-month	12-month		3-month	6-month	9-month	12-month
GE	0.946	0.947	0.949	0.950	US	0.852	0.861	0.935	0.844
UK	0.553	0.570	0.629	0.585	UK	0.725	0.721	0.821	0.810
CA	0.637	0.641	0.560	0.644	CA	0.698	0.713	0.687	0.759
FR	0.507	0.517	0.257	0.892	FR	0.862	0.867	0.940	0.527
JA	0.836	0.834	0.837	0.784	JA	0.287	0.295	0.262	0.325
NO	0.529	0.582	0.633	0.659	NO	0.587	0.603	0.529	0.575
SW	0.890	0.897	0.881	0.875	SW	0.527	0.505	0.350	0.256

Notes. The tables present estimation results from a state space model, where the interest rate differentials of seven countries with respect to the US and Germany are decomposed into a global factor component and an idiosyncratic component, as specified in Section 4.2.1. *Panel A)* presents the global factor loadings (estimates of the A vector) with standard errors in parentheses. *Panel B)* displays the percentage of the interest rate differential variance that is explained by the common factor, for each of the four maturities m. One, two and three asterisks denote statistical significance at the 10, 5 and 1 percent level respectively.

Table 4.2 Correlation between global factors and policy rates

	3-month		6-month		9-month		12-month	
	Factor US	Factor Ge	Factor US	Factor Ge	Factor US	Factor Ge	Factor US	Factor Ge
US	0.902	0.740	0.900	0.738	0.903	0.815	0.865	0.719
GE	0.250	0.073	0.260	0.075	0.281	0.119	0.223	0.063

Notes. The tables present the correlation coefficients between the global factor of the interest rate differentials with respect to the US (Factor US) and Germany (Factor Ge) with the US (US) and German (GE) interest rates for the each maturity m.

Table 4.3 Monetary policy and the global factor (OLS specification)

Panel A) 3-month data ($m = 3$)

Model	EU		US		EU-US	
Variables	$M_{1,EU}$	$M_{2,EU}$	$M_{1,EU}$	$M_{2,US}$	$M_{1,EU/US}$	$M_{2,EU/US}$
Constant	0.826*** (0.242)	0.735*** (0.371)	1.703*** (0.125)	1.896*** (0.214)	1.059*** (0.125)	0.988*** (0.223)
FX_t^6	1.118*** (0.224)	1.112*** (0.244)	0.405*** (0.154)	0.370*** (0.163)	0.198 (0.125)	0.116 (0.121)
$Eonia$	−0.219*** (0.064)	−0.229*** (0.096)			0.397*** (0.046)	0.580*** (0.067)
$inflation_{EU}$		11.842 (17.705)				−16.323* (9.133)
$output\_gap_{EU}$		−3.458 (7.116)				10.951* (6.898)
FF			−0.415*** (0.029)	−0.409*** (0.034)	−0.608*** (0.032)	−0.685*** (0.037)
$inflation_{US}$				−3.763 (8.807)		−8.505 (7.347)
$output\_gap_{US}$				−5.754 (0.541)		3.741 (4.662)
Adj. R^2	0.231	0.200	0.681	0.678	0.795	0.842

(continued...)

(...Table 4.3 continued)

Panel B) 6-month data ($m = 6$)

Model	EU		US		EU-US	
Variables	$M_{1,EU}$	$M_{2,EU}$	$M_{1,EU}$	$M_{2,US}$	$M_{1,EU/US}$	$M_{2,EU/US}$
Constant	0.744*** (0.215)	0.573*** (0.321)	1.466*** (0.128)	1.492*** (0.227)	0.932*** (0.133)	0.872*** (0.240)
FX_t^6	1.184*** (0.136)	1.192*** (0.146)	0.610*** (0.112)	0.617*** (0.122)	0.396*** (0.100)	0.222** (0.104)
$Eonia$	−0.179*** (0.057)	−0.194** (0.087)			0.356*** (0.051)	0.585*** (0.077)
$inflation_{EU}$		19.187 (15.384)				−16.575* (9.612)
$output\_gap_{EU}$		−3.808 (6.231)				10.238* (8.253)
FF			−0.346*** (0.030)	−0.337*** (0.035)	−0.535*** (0.037)	−0.642*** (0.043)
$inflation_{US}$				3.934 (9.097)		−9.923 (8.081)
$output\_gap_{US}$				−7.089 (5.323)		3.335 (5.001)
Adj. R^2	0.425	0.411	0.700	0.695	0.783	0.828

(continued...)

(...Table 4.3 continued)

Panel C) 9-month data ($m = 9$)

Model	EU		US		EU-US	
Variables	$M_{1,EU}$	$M_{2,EU}$	$M_{1,EU}$	$M_{2,US}$	$M_{1,EU/US}$	$M_{2,EU/US}$
Constant	0.744*** (0.183)	0.745*** (0.268)	1.227*** (0.120)	0.911*** (0.215)	0.874*** (0.133)	0.732*** (0.243)
FX_t^9	1.187*** (0.092)	1.178*** (0.099)	0.771*** (0.086)	0.829*** (0.093)	0.617*** (0.086)	0.512*** (0.099)
$Eonia$	−0.166*** (0.048)	−0.216*** (0.076)			0.249*** (0.052)	0.446*** (0.091)
$\underset{EU}{inflation}$		10.180 (13.225)				−22.542** (9.925)
$\underset{EU}{output\_gap}$		1.671 (5.216)				10.022** (4.357)
FF			−0.276*** (0.028)	−0.265*** (0.032)	−0.416*** (0.039)	−0.518*** (0.052)
$\underset{US}{inflation}$				18.820** (8.786)		2.407 (9.159)
$\underset{US}{output\_gap}$				−7.661 (4.837)		0.720 (5.513)
Adj. R^2	0.615	0.606	0.762	0.765	0.798	0.828

(continued...)

(...Table 4.3 continued)

Panel D) 12-month data ($m = 12$)

Model	EU		US		EU-US	
Variables	$M_{1,EU}$	$M_{2,EU}$	$M_{1,EU}$	$M_{2,US}$	$M_{1,EU/US}$	$M_{2,EU/US}$
Constant	0.740*** (0.169)	1.139*** (0.239)	1.131*** (0.111)	0.859*** (0.192)	0.840*** (0.127)	0.811*** (0.217)
FX_t^{12}	1.034*** (0.068)	1.048*** (0.073)	0.726*** (0.065)	0.755*** (0.069)	0.625*** (0.066)	0.608*** (0.072)
$Eonia$	−0.159*** (0.044)	−0.280*** (0.069)			0.199*** (0.050)	0.334*** (0.085)
$inflation_{EU}$		−10.336 (12.023)				−34.702*** (9.243)
$output\_gap_{EU}$		10.674 (4.576)				12.638*** (3.866)
FF			−0.246*** (0.026)	−0.263*** (0.030)	−0.357*** (0.037)	−0.446*** (0.046)
$inflation_{US}$				12.374 (7.941)		7.533 (8.088)
$output\_gap_{US}$				1.154 (4.461)		2.506 (4.918)
Adj. R^2	0.694	0.806	0.805	0.660	0.827	0.864

Notes: The tables show the results from an OLS regression of the common factor of interest rate differentials (dependent variable) on various monetary policy variables (independent variables). The different variables presented in the first column relate to the independent variables of models $M_{1,c}$ or $M_{2,c}$ for c=US, EU, US/EU, as specified in Section 4.2.2. Different panels correspond to different maturities (m) for the interest rate differentials, where $m = 3, 6, 9$ and 12-months. The tables report coefficient estimates and standard errors (in parentheses). They also display the adjusted R-squared measure of overall fit (Adj. R^2). One, two and three asterisks denote statistical significance at the 10, 5 and 1 percent level respectively.

Table 4.4 Monetary policy and the global factor (IV specification, 3-month data, m = 3)

Model	EU		US		EU-US	
Variables	$M_{1,EU}$	$M_{2,EU}$	$M_{1,EU}$	$M_{2,US}$	$M_{1,EU/US}$	$M_{2,EU/US}$
Constant	0.769*** (0.262)	0.243 (0.592)	1.630*** (0.127)	1.841*** (0.216)	1.046*** (0.128)	0.996*** (0.225)
FX_t^3	1.133*** (0.229)	1.152*** (0.283)	0.436*** (0.156)	0.401** (0.164)	0.241* (0.127)	0.120 (0.120)
$Eonia$	−0.202*** (0.071)	−0.051 (0.120)			0.375*** (0.048)	0.566*** (0.069)
$inflation_{EU}$		40.258 (39.756)				−16.572* (9.039)
$output\_gap_{EU}$		−39.545*** (13.832)				11.599* (10.865)
FF			−0.397*** (0.029)	−0.387*** (0.034)	−0.585*** (0.033)	−0.667*** (0.037)
$inflation_{US}$				−3.636 (8.873)		−7.890 (7.263)
$output\_gap_{US}$				−7.024 (5.476)		1.948 (4.643)
$J-stat.$	0.061	0.051	0.108	0.057	0.268	0.999

Notes: The table shows the results from a 2SLS regression of the common factor of 3-month interest rate differentials (dependent variable) on various monetary policy variables (independent variables). Lagged values of the independent variables are being used as instruments. The different variables presented in the first column relate to the independent variables of models $M_{1,c}$ or $M_{2,c}$ for c=US, EU, US/EU, as specified in Section 4.2.2. The table reports coefficient estimates and standard errors (in parentheses). It also displays the results from Sargan's test for overidentifying restrictions ($J-stat.$). One, two and three asterisks denote statistical significance at the 10, 5 and 1 percent level respectively.

Chapter 5

5 Conclusion

In this thesis we investigate three different topics in international empirical finance, which, despite the considerable amount of attention focused on them in previous years, are still currently under debate.

In Chapter Two, armed with several tests proposed by the literature for testing forward rate unbiasedness in the FX market, we provide a simple, yet intuitive bridge to a different derivatives market, the currency options market, as a vehicle of forming expectations about future spot exchange rates. Our main focus is on performing tests of the unbiasedness hypothesis. To that end, we use the conventional forward rate and also introduced an option equivalent (synthetic forward) contract. We then apply some prominent tests of unbiasedness, based on the standard UIP condition in a stationary setting as well as cointegration tests for unbiasedness of the (synthetic) forward rate, the latter combined with residual autocorrelation tests.

This research provides encouraging results. We manage to bridge the distance between the forward (OTC) market and the options (exchange traded) market, by directly comparing the test results obtained for the two markets. Viewed from a different angle, our research offers a novel robustness check to the tenacity of the well-documented forward bias anomaly that characterises the relevant literature.

We record no qualitative difference between the two types of derivatives products in our results. Specifically, our results suggest the existence of an "options bias," similar to the forward bias, frequently recorded in the relevant literature estimating stationary regressions of the exchange rate change on the lagged forward premium. This finding indicates, in turn, violation of market efficiency in its risk neutral formulation as implied by UIP, possibly as a consequence of the existence of a risk premium or possibly because of the estimation problems induced by the persistence of the forward (option) premium and of exchange rate volatility. We therefore shift our attention to a cointegration and a residual correlation analysis that allows for the endogeneity problems caused by a potential unobserved risk premium term and requires less stringent conditions on the error term process. This analysis attests that indeed the (synthetic) forward is an unbiased predictor of the future spot rate. Overall, we interpret the evidence in this exercise as suggesting that forward and options provide optimal exchange rate predictions consistent with the notion of unbiasedness.

Chapter Three elaborates on the long standing PPP puzzles. Earlier literature seeked answers by employing unit root tests with different levels of sophistication. Amongst those, the ones which accommodated the non-Gaussian behaviour of the RER, seemed to have better power in detecting reversion towards the PPP equilibrium. In this chapter, we present QAR semi-parametric and non-parametric methods as an alternative approach for robust inference in non-gaussian series. The quantile approach adopts an agnostic and yet flexible framework for the analysis of the RER behaviour, thus sidestepping the need to specify theory-consistent driving forces of the RER dynamic adjustment process. More precisely, the quantile framework makes no assumptions about the underlying distribution of the RER, while allowing for different (symmetric or asymmetric) persistence patterns at the different quantiles. It this sense, it is possible to observe sequences of unit-root behaviour, while occasional mean reverting tendencies can undermine the persistence of the whole process. By taking into account the different adjustment processes at the different quantiles, the quantile approach offers a more robust unit root test than standard alternatives.

More importantly, the QAR analysis and inference sheds light into both PPP puzzles. As concerns the first one, our methodology offers some support for the PPP, by providing evidence in favour of a mean reversion in the RER from two different quantile unit root tests. Our approach also addresses the second PPP puzzle by undertaking a detailed analysis of the impact of different magnitudes of actual shocks on the RER. We rationalise the high persistence of the RER behaviour, by suggesting that different magnitudes of shocks can induce different speeds of adjustment to the RER, while maintaining consistency to the limit to arbitrage theory.

More specifically, our evidence from two different quantile unit root tests in semi- and non-parametric settings suggests that the RER is not a constant unit root process across quantiles. We find that the bigger the shock to the RER (i.e. the bigger the quantile) the faster the mean reversion back towards its long run equilibrium, with half lives comfortably less than a year, in the case of extreme shocks. Our results are further enhanced when large shocks hit the RER at points already far from its equilibrium. In such cases half lives can fall significantly less than a year. However, the mean reversion ability of large shocks is diminished in cases when the RER is around its equilibrium value. Finally, in the absence of shocks, mean reversion cannot be established irrespective of the RER disequilibrium level. In addition, our method captures asymmetric dynamic adjustment of the RER, i.e. positive shocks have different impact than negative shocks. Our results offer novel insights on the RER mean reverting behaviour and further refine and enhance previous evidence in the PPP literature.

Finally, Chapter Four of this thesis aims to analyse the international linkages in the money market rate differentials with respect to the US and Germany across a group of countries. We interpret these differentials as reflecting the monetary policy deviations of each country versus the denominating country (US or Germany). We employ a maximum likelihood, dynamic factor model where the interest rate differentials of a group of countries vis-á-vis a common denominating country are modeled as the sum of a common/global component and an idiosyncratic/domestic component. In that sense, the global factor captures the common fluctuations in the policy responses of our group of countries versus the denominating country. We provide an analysis of the relative importance of global versus domestic factors and we also try to identify the special role of the two major economies, the US and the EU in driving the global factor itself.

The main contribution of our study relies on the use of factor analysis methods on short term interest rate differentials. Adding to previous studies, our study provides a measure of comovements in the monetary policy reactions across countries and the importance of these comovements on the policy fluctuations in each of the countries under investigation. We also get an indication of the existence of dominant countries, either in the sense that they affect the strength of the global factor or that their policy fluctuations drive the global factor fluctuations. We then further investigate finer interactions amongst our players, and specifically whether the policy deviations of a group of countries from the denominating country are explained by the pattern set by the deviations of another dominant country. We also explore whether these dominant countries can set their policies independently of others and yet command other countries' responses. Finally, we provide indications for the role of the US as the global interest rate.

We reveal valuable insights in the behaviour of the short-term interest rate differentials with direct implications for monetary policy actions. We find that there is a significant common global component present in the fluctuations of the interest rate differentials for almost all countries, although the patterns of interaction seem to be different for the US and Germany, revealing different spheres of US and German policy influence and different degrees of regional integration. For example, EU country differentials are much more integrated vis-á-vis Germany than vis-á-vis the US. The opposite is happening for Japan and Switzerland. Based on this result, we can further argue that the US and the EU are in a position to choose their policies independently, although the interaction of their policies seems to be followed by the rest of the world. Finally, our analysis suggests that the US interest rate is the dominant force behind the global factor, thereby providing support for the role of the US as the global interest rate.

Overall, our analysis accepts that the monetary policies between a set of countries may be different. However, we find that a large part of these differences are in fact responding to common considerations and are shared across countries. More importantly, they tend to follow the policy interactions of two major global players, namely the US and the EU. Therefore, the policy implication would be that monetary policy-makers should pay closer attention to US and EU policies when forming their own.

This thesis, although not claiming to have analysed all possible aspects of the topics treated therein, enjoys some success at further refining and enhancing already existing paths of literature, while shedding light to new, potentially promising strands or research.

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